Common stochastic trends and policy shocks in the open economy

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Publication date:
1997

Document Version
Publisher's PDF, also known as Version of record

CentER
for
Economic Research

Common Stochastic Trends and Policy Shocks in the Open Economy:

Empirical Essays in International Finance and Monetary Policy

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Proefschrift ter verkrijging van de graad van doctor aan de Katholieke Universiteit Brabant, op gezag van de rector magnificus, prof. dr. L.F.W. de Klerk, in het openbaar te verdedigen ten overstaan van een door het college van decanen aangewezen commissie in de aula van de Universiteit op

maandag 26 mei 1997 om 14.15 uur door

Francis Yao Kumah

geboren op 24 augustus 1958 te Accra, Ghana.

Promotor: Prof. dr. Harald Uhlig

To Leticia, Selorm and Cindy

"The only way .... our science can give positive advice on a large scale to politicians and businessmen, leads through quantitative work. For as long as we are unable to put our arguments into figures, the voice of our science, although occasionally may help dispel gross errors, will never be heard by practical men. They are, by instinct, econometricians all of them, in their distrust of anything not amenable to exact proof."

Joseph A. Shumpeter(1933),

"The Common Sense of Econometrics", *Econometrica* 1, page 12.

Contents

List of Figures

List of Tables

Preface

This dissertation was made possible by financial assistance received from Tilburg University on the basis of a contract as an Assistent i Opleiding (AIO) - i.e. as a teaching assistant - at CentER for Economic Research. For this I am very grateful to Tilburg university. After only a couple of months of my stay at CentER for Economic Research - at Tilburg University - it dawned on me that I had finally realized my dream of working with enthusiastic researchers and supporting staff alike in a *scientific* research institute. I owe a great debt of gratitude, first and foremost, to my thesis supervisor, Prof. Harald Uhlig for very insightful discussions and comments on the various chapters of this dissertation. I am very grateful to Prof. Herman Bierens as well for helpful comments on the econometric aspects of a couple of chapters of this dissertation. I also owe a great debt of gratitude to Professors Auther van Soest, Theo Nijman and Lucrezia Reichlin for their contributions as members of my Ph. D. defense committee.

Chapter 5 of this dissertation was presented at a conference organised by Centre for Economic Policy Research (CEPR) - *Model Specification, Identification and Estimation in Empirical Macroeconomics* - in Perugia, Italy, 9th-12th January, 1997 at the invitation of Lucrezia Reichlin. Thanks, Lucrezia, for inviting me; it was a very enlightening experience to meet and discuss my work with all those seasoned economists. I am very grateful to all the participants at the conference - especially to Prof. Fabio Canova at Università di Modena in Italy - for all the valuable comments on my presentation at the conference.

I also thank Shell for granting me a travel stipend to enable me present aspects of my research at the Finance Division of the 25th Annual Conference of the Administrative Sciences Association of Canada (ASAC) at St. John's, Newfoundland, in May, 1997.

I would also like to offer sincere thanks to all my colleagues in the Ph. D. programme at Tilburg University. My utmost thanks go to my office-mates Anurag Banerjee and Jan Fidrmuc for creating such a friendly and research-stimulating atmosphere at our office. Others who

have contributed in making my period of research sojourn here at CentER pleasant include Jimmy Miller (the general all-round nice bloke), Jan Lemmen, Erik Canton, Trea Laske, Bas van Aarle, Marco Hoeberichts, Henri de Groot, Mohammed Nabih, Zafar Iqbal and all others. I deeply cherish the nice times and discussions we had during the lunch seminars in macroeconomics. Also, my colleague Salifu Ibrahim deserves lots of thanks for helping me to gain a deeper insight into some of the issues that arose while we were co-authoring a couple of papers included in this dissertation. To the secretaries at CentER who work conscientiously to create an atmosphere very conducive to research I say "*hartelijk bedankt*". In my mother tongue this would be put simply as "*akpe*" - meaning thanks. I would surely miss the tea-times.

Before ending this preface I would like to take this opportunity to express sincere thanks to my dear wife - Leticia F. Ayi - for being very understanding, creating a peaceful atmosphere at home and thereby making it possible for me to work both at school and at home. Thanks to my kids too for co-operating even when they were denied the right to play with daddy because he had to think of some econometric models and computer programmes.

Tilburg, March, 1997 Francis Y. Kumah

Chapter 1:

Introduction and Summary

1.1 Background

The demise of the Bretton Woods system of pegged exchange rates in the early 1970s gave birth to a lot of theoretical and empirical research aimed at explaining the observed volatility of exchange rates and fluctuations in the current accounts of nations. Also, the early 1980s witnessed the world economy plagued by large and uncoordinated innovations in monetary and fiscal variables as well as by high and volatile real interest rates and exchange rate misalignments - i.e. the tendency for the real exchange rate movements to diverge from the long run equilibrium path as measured by relative purchasing power parity. The increased number of articles in exchange rate economics as an aspect of international finance can only be explained by the juxtaposition of the events listed above. Prominent among the research efforts in this area are the monetary models - on which modern target zone models of Krugman(1991) and others are based - and portfolio-balance models of exchange rate determination. In a highly influential paper Mussa(1986) - in the same vein as Dornbusch(1976) - assigns a role to sluggish price adjustments in explaining the observed exchange rate volatility. Works by Branson - see for instance Branson(1984) or Branson and Hendersen(1985) - fall mainly within the category of portfolio-balance models seeking the explanation among factors bordering on capital mobility, fiscal and monetary factors as well as productivity shifts. Stockman(1987) argues that the observed volatility of exchange rates following the collapse of the Bretton Woods system could be better explained by real productivity shocks rather than by price rigidities. Empirical findings by Huizinga(1987) - that most real exchange rate time series do appear to possess a unit root and that the bulk of the forecast error variance of real exchange rates are attributable to permanent real shocks - do support this view. De Grauwe(1989) concludes however that though, within some sub-periods, real exchange rate changes coincide with real shocks "real exchange rate movements of the last fifteen years cannot easily be explained by real shocks" - (see page 122). Though MacDonald and Taylor(1994), using modern techniques of cointegration, find evidence supporting the use of the monetary model of exchange rate determination, the empirical performance of the asset market based theories of exchange rate determination generally leaves much to be desired. Thus, despite the apparent heavily-researched nature of exchange rate economics most of the empirical findings do not seem to support the conclusions of the underlying theoretical models. There is therefore still room for further research not only on the basic issue of determining the relative significance of the real and nominal factors underlying exchange rate fluctuations but also in solving the puzzles generated by the deviation of observations on the market from their theoretical predictions.

This dissertation contributes to empirical international finance by introducing and using new empirical methods in identifying productivity shocks and monetary policy shocks. The shocks so identified are then used in explaining observed exchange rate and current account fluctuations. The rationale of the studies conducted is to blend international monetary theory with modern time series analysis in an attempt to contribute to our understanding of aspects of the international monetary system. These aspects include the effects of productivity and policy shocks on capital flows and bilateral exchange rate fluctuations. The international business cycles literature has some contributions involving the effects of productivity shocks on investment and current account fluctuations - see for instance Baxter and Crucini(1993), Backus, Kehoe and Kydland(1992) and Mendoza(1991) for the main ingredients of this approach. The measurement of the Solow parameter under the aforementioned approach relies on the assumptions (of perfect competition, constant returns to scale and very insignificant variations in the capital input over the cycle) that yield an overestimation of the variance of the productivity/technology shocks. To circumvent these shortcomings chapters 3 and 4 in the first part of this dissertation adopt the King-Plosser-Stock-Watson common stochastic trends approach in analysing fluctuations in investments, current accounts and bilateral exchange rates in response to shocks to stochastic trends. In chapter 3 these stochastic trends are identified and characterized, using a priori theoretical reasoning, as real (productivity) trends or nominal trends. In chapter 4 the distinction is between idiosyncratic (country-specific) and global productivity shocks.

Many central bankers - including Alan Greenspan (chairman of the Federal Reserve of the United States), Governor Thiesen (of the Central Bank of Canada) and Hans Tietmeyer (president of Germany's Bundesbank) - believe that a cut in the budget deficit would enhance the external value of the domestic currency even though international macroeconomists are not unanimous on the direction of the effects of such budget deficit cuts. In fact in a speech at a symposium organized by the federal reserve of Kansas City, the Chairman of the federal reserve system, Alan Greenspan is quoted as saying "... the point the central bankers are making is that lower long-term inflation expectations can significantly overwhelm the short-term interest rate effects, (*of reductions in the budget deficit of the United States*) and through arbitrage back to the spot rate,..." (see Greenspan(1995), page 141, *italics mine*). However, within the context of the general asset market models of exchange rate determination, changes in the domestic budget deficit are inextricably associated with changes in the country's net foreign assets position, in perceived riskiness of domestic assets (which influences the size of the risk premium desired by investors for continued holding of domestic assets) and in expectations regarding the future value of the domestic currency[1]. Thus changes in the budget deficit have both short run as well as long run implications for the external value of the domestic currency. In the short run, an increase in the budget deficit causes a sudden appreciation of the domestic currency (as the domestic short-term interest rate rises) creating a trade deficit and worsening the current account (as foreign capital flows into the country). Capital inflows continue until the risk premium (as the increased budget deficit increases

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[1] The arguments of the competing model - the Mundell-Flemming model - is that an increase in the budget deficit increases domestic spending. Given that the supply of money is held constant, the argument goes, the increase in spending instigates domestic short-term interest rate increases as the excess demand for goods is translated into increased demand for money. The increase in the short-term interest rate encourages capital inflows that strengthen the domestic currency vis-a-vis foreign currencies and generate trade deficits.

investors' perceived riskiness of domestic assets) associated with domestic assets increase and/or expectations of future domestic currency depreciation increase sufficiently to equalize expected returns on domestic and foreign assets. Thus, uncovered interest parity must hold in the long run suggesting that positive fiscal innovations must necessarily be associated with long run depreciation of the domestic currency. These are the issues that chapter 6 of this dissertation investigates using the empirical tools of cointegration and vector-autoregressions.

1.2 Outline of the Study

The objective of this introductory chapter is to present the rationale of the whole thesis as well as to give a brief summary of the main findings of the chapters of the dissertation. A brief note on the modern time series methods of common stochastic trends is presented here as well. Chapter 2 presents a brief overview of the theoretical models underlying the discussions in subsequent chapters. The dissertation is divided into two self-contained parts with two essays in each part. The essays in each of these parts are slightly revised versions of published papers and discussion papers. The first set of essays presented in this dissertation use the common trends econometrics approach extensively in addressing, among other issues, exchange rate and current account fluctuations. The second part takes up the inconclusive empirical issue of the effects of monetary and fiscal policy shocks on exchange rate and current account fluctuations. The main concerns of each of these parts of the dissertation are summarized below in subsections 1.2.1 and 1.2.2 respectively.

1.2.1 Common Stochastic Trends in the Open Economy

This section reviews the common trends statistical model used to characterize the nature of shocks and the responses of our variables of interest in the first part of the dissertation.

Consider any n x 1 dimensional macroeconomic time series X_t characterised by the vector moving average representation

$$(1 - L)\, X_t \;=\; \delta + C(L)\varepsilon_t \tag{1.1}$$

where $C(L) = I_n + C_1 L + C_2 L^2 + \ldots$ (an invertible lag matrix polynomial), ε_t is white-noise with a zero mean vector (i.e. $E[\varepsilon_t] = 0$) and a positive definite covariance matrix, $\Xi = E[\varepsilon_t \varepsilon_t']$, and L is the lag operator such $L^i \chi_t = \chi_{t-i}$. Under fairly general conditions it is possible (using matrix algebra) to find another matrix polynomial $C^*(L)$ such that

$$C(L) = C(1) + (1 - L)C^*(L) \tag{1.2}$$

where the polynomial matrix $C^*(L)$ is given by

$$\begin{aligned} C^*(L) &= (1 - L)^{-1}(C(L) - C(1)) \\ &= \sum_{j=0}^{\infty}\left(\sum_{i=0}^{j}(C_i - C(1)\right)L^j = \sum_{j=0}^{\infty}\left(\sum_{i=j+1}^{\infty} C_i\right)L^j = -\sum_{j=0}^{\infty} C_j^* L^j \, . \end{aligned}$$

Utilizing equation (1.2) and the usual convention of letting ε_s be zero for s ≤ 0, and with X_0 representing the non-random initial value of X_t, then by recursive substitution[2] of equation (1.1) we obtain,

$$X_t = X_0 + \delta t + C(1)(1 + L + L^2 + \ldots + L^{t-1})\varepsilon_t + C^*(L)\varepsilon_t \tag{1.3}$$

[2] See Stock and Watson(1988) for the procedure applied here.

Since δ is one of the vectors in the nullspace of $C(1)$, it can be expressed as a linear combination of the columns of $C(1)$, say $\delta = C(1)\upsilon$ where υ is an $n\times 1$ vector. Substituting this expression into equation (1.3) above we obtain

$$X_t = X_0 + C(1)[t\upsilon + \sum_{i=0}^{t}\varepsilon_i] + C^*(L)\varepsilon_t$$

Denoting the expression in the square brackets (the random-walk component) by $\tau_t = \mu + \tau_{t-1} + \varphi_t$, we obtain, after some manipulations, the expression

$$\begin{aligned} X_t &= X_0 + A\tau_t + C^*(L)\varepsilon_t \\ &= X_0 + A[\mu + \tau_{t-1} + \varphi_t] + C^*(L)\varepsilon_t \end{aligned} \qquad (1.4)$$

where $C(1)\varepsilon_t = A\varphi_t$, $C(1)\delta = A\mu$, and hence (for $E[\varphi_t\varphi_t'] = \Phi$) $C(1)\Xi C(1)' = A\Phi A'$. Assuming, for ease of interpretation, that $E[\varphi_t] = 0$ **and** $E[\varphi_t\varphi_t'] = \Phi = I_{n-r}$, we see that $C(1)\Xi C(1)' = A\Phi A' = AA'$. Further, pre-multiplying equation (1.4) by β' (where β is the cointegrating vector of X_t) we see that for the structure of the A matrix to be consistent with the cointegration requirement (of stationarity of $\beta' X_t$) implies the requirement that $\beta' A = 0$. Equation (1.4) is the multivariate version of the Beveridge-Nelson decomposition[3] of the cointegrated vector moving average representation, (1.1), expressing X_t as a linear combination of $n - r$ linearly independent stochastic trends (the common trends that have permanent effects on X_t) and transitory components, $C^*(L)\varepsilon_t$ which are stationary. X_0 contains the initial values of X_t.

$$\tau_t = \mu + \tau_{t-1} + \varphi_t \qquad (1.5)$$

[3] See Beveridge and Nelson (1981) and/or Stock and Watson (1988) for the details of this decomposition.

is a $k \times 1$ vector of non-stationary variables (with structural white-noise shock vector, φ_t, whose variance-covariance matrix is denoted by Φ) that drive the system and μ is the drift term. A is an $n \times k$ matrix of coefficients to be characterized (by the aid of the underlying theory) and estimated. It measures the long-run impact of the k common stochastic trends in τ_t on each of the elements of the X_t matrix. The transitory part of the model is described by the $C^*(L)$ matrix polynomial.

Impulse Responses and Variance Decompositions

On the basis of the analysis of Campbell and Shiller(1988), Warne(1990) shows that in the case of cointegration the impulse response functions and variance decompositions can be obtained by inverting a particular restricted VAR model of the form

$$H(L)y_t = \delta^* + \epsilon_t^* \tag{1.6}$$

To show the relationship between the variables and the parameters of the error correction (EC) model and this RVAR define the following matrices:

$$M^* = \begin{bmatrix} S_k \\ \beta' \end{bmatrix}, \quad \alpha^* = [\ 0 \ \ \alpha\ ], \quad D^*(L) = \begin{bmatrix} I_k & 0 \\ 0 & (1-L)I_r \end{bmatrix}, \quad and \quad D_{\perp}^*(L) = \begin{bmatrix} (1-L)I_k & 0 \\ 0 & I_r \end{bmatrix}$$

where S_k is a $k \times n$ matrix chosen such that its rows are linearly independent of those of b', the rank of M^* is n and α^* is an nxn matrix. Given these definitions we can further derive the following relationships

$$\delta + \alpha\beta' X_{t-1} + \epsilon_t = D(L)D^*(L)D_{\perp}^*(L)X_t$$
$$= D(L)M^{*\,-1}D^*(L)D_{\perp}^*(L)M^*X_t \tag{1.7}$$

Further, define an $n\times 1$ dimensional $I(0)$ matrix of time series vectors denoted y, such that $y_t = D_{\perp}^*(L)M^*X_t$ so that we can re-write equation (1.7) as $\delta + \alpha\beta' X_{t-1} + \varepsilon_t = D(L)M^{*-1}D^*(L)y_t$. Thus, for $H(L) = M^*[D(L)M^{*-1}D^*(L) + \alpha^*L]$ where $H(0) = M^{*-1}$ (and hence $M^*H(0) = I_n$) we obtain the RVAR model given that $\alpha\beta' X_{t-1} = [0\ \alpha][\Delta S_k X_{t-1}\ \beta' X_{t-1}]'$, $\delta^* = M^*\delta$ and $\varepsilon^*_t = M^*\varepsilon^*_t$. Once we estimate the above equation - (1.6) - we can also easily estimate the Wold moving average representation (1.1) and hence the impulse response functions as well as the variance decompositions. However to calculate the impulse responses and the variance decompositions associated with the shocks to the common trends we need some additional restrictions on the model so far specified.

Let $\eta_t = [\varphi_t',\ \psi_t']' = F\varepsilon_t$ be a matrix of shocks where $F = [F_k',\ F_r']'$, φ_t may be considered as an n-dimensional vector of k permanent shocks and ψ_t as an $r\times 1$ vector of transitory shocks. The permanent shocks, φ_t, and the transitory shocks, ψ_t, are assumed independent. Further, define $R(\lambda) = C(\lambda)F^{1}$. Then given the specification/definitions under (1.6) above we observe that $\eta_t = (A'A)^{-1}A'C(1)\varepsilon_t = F\varepsilon_t$ and hence the Wold vector moving average representation can be re-written as

$$(1 - L)X_t = \delta + R(L)\eta_t \tag{1.8}$$

where $\Gamma = E[\eta_t\eta_t']$ is a diagonal covariance matrix that can be partitioned as

$$\Gamma = \begin{bmatrix} \Phi & 0 \\ 0 & \Psi \end{bmatrix}$$

$\Psi = E[\psi_t \psi_t']$ is diagonal matrix representing the covariance matrix of the transitory shocks. Given the F_k matrix as defined above, Warne(1990) has shown that $F_r = Q_r^{-1} \alpha \Xi^{-1}$ where Q_r is an $r \times r$ invertible matrix satisfying the assumption of the independence of permanent and transitory shocks (ie. $E[\varphi_t \psi_t'] = (A'A)^{-1} A'C(1) \Xi F_r' = 0$ since, as shown by Engle and Granger(1987), $C(1)\alpha = 0$) and is chosen such that the covariance matrix of the transitory shocks is diagonal. In practice the identification of the transitory shocks (ψ_t) requires $r(r - 1)/2$ zero-restrictions on their contemporaneous effects on the endogenous variables, X_t.

The expression $R(L)\eta_t$ in equation (1.8) above represents the impulse response function of ΔX_t. This implies that if we shock ΔX_t by a one standard deviation change in η_t the dynamic responses of ΔX_t s periods after the shock can be expressed as $resp(\Delta X_{t+s}) = R_s$. To estimate these responses, all we need are the estimates of C_s and F^1 since the estimate of R_s is given by $R_s = C_s F^1$. Similarly forecast error variance decompositions could be obtained as well. In fact, Warne(1990) has shown that the s steps ahead forecast error covariance matrix can be expressed as

$$\begin{aligned} V_s &= \sum_{j=1}^{s} \left(\sum_{j=1}^{s+1-m} R_{j-1} \right) \left(\sum_{j=1}^{s+1-m} R_{j-1}' \right) \\ &= \sum_{m=1}^{s} R_m^* R_m^{*\prime} \end{aligned} \tag{1.9}$$

For a finite s the variance decompositions V^*_{il},s follows directly from equation (1.9) above where il denotes the (i,l):th element of R_{j-1}. However, since $\lim_{s\to\infty} V_s$ is not infinite, Warne shows again that equation (1.9) can be re-written (by collecting products $R_{j-1}R'_{j-1}$ into one group and all others into another group) as

$$V_s = \sum_{j=1}^{s}(s + 1 - j)R_{j-1}R_{j-1}' + \sum_{m=1}^{s-1}\sum_{j=1}^{m}(s - m)\left(R_{j-1}R_m' + R_m R_{j-m}'\right) \tag{1.10}$$

Defining V_∞ as $\lim_{s\to\infty} V_s/s$ we derive from equation (1.10) above the expression

$$\begin{aligned} V_\infty &= \sum_{j=1}^{\infty}R_{j-1}R_{j-1}' + \sum_{m=1}^{\infty}\sum_{j=1}^{m}\left(R_{j-1}R_m' + R_m R_{j-1}'\right) \\ &= \left(\sum_{j=1}^{\infty}R_{j-1}\right)\left(\sum_{j=1}^{\infty}R_{j-1}'\right) \\ &= R(1)R(1)' \end{aligned} \tag{1.11}$$

which can then be used in obtaining the s steps ahead forecast error variance decomposition for the n variables in our X_t vector.

Empirical Applications

Apart from King et al(1991) who apply the common trends methodology as outlined above in analysing fluctuations in macroeconomic time series within the context of a closed economy, we are not aware of any research integrating the nonstructural framework of the common trends approach with theoretical structures. Within a non-stochastic framework however, Ahmed(1987) analyses fluctuations in the balance of trade of the UK treating real military spending as the main forcing variable and then models it as a random walk with drift. King et al(1991) utilize a simple stochastic growth model where consumption, investments and output have a common trend arising from the technology factor of an aggregate Cobb-Douglas production function. Basing their arguments essentially on a priori reasoning, rather than on rigorous theoretical or empirical reasoning, they assume there are two stochastic trends in their set of variables extended to include money and prices. King et al(1991) assume that the logarithm of the velocity of money, the logarithms of the ratios of consumption and output as well as of investments and output are stationary - leaving them with two non-stationarities. The approach, as we apply it in the first part of this dissertation, bases all required restrictions for identifying (and distinguishing among the various stochastic trends) on theoretical reasoning mainly from international monetary theory and policy. This is the main feature that distinguishes our application of the common trends approach from that of King et al(1991) - that is if we ignore the obvious difference of ours being within an open economy context whereas theirs is a closed economy analysis.

The first part of this dissertation applies the common trends approach, as outlined above, in answering open economy questions on the relative effects of real and nominal trends on exchange rate and the current account fluctuations. The first chapter of this first part - chapter 3 - investigates the effects of real and nominal trends on interest rate differentials (between the US and each of the four countries Germany, Japan, Sweden and the UK), exchange rates (defined here as units of domestic currency per US dollar), and the current account using quarterly time series data on output, budgetary deficits, money supply, interest rate differentials exchange rates and the current account balance for these countries over the postwar period 1974:1 - 1994:4. We found evidence that domestic technological trends (or supply shocks) do have more significant effects on the exchange rate and current account balance over the short and the long run than nominal trends do. Our results also show, not unexpectedly, that transitory shocks affect the exchange rate and the current account only *transitorily* (that is to say their effects are felt mainly in the short and medium term).

The second chapter in the first part of the dissertation also applies the common trends approach in analysing current account fluctuations. The chapter uses the intertemporal approach to current account fluctuations (see for instance Obstfeld(1986), Razin(1992) and/or Glick and Rogoff(1995)) as its theoretical backbone. In this chapter we show that rather than using the Solow parameter as a proxy for productivity shocks one could utilize the common trends approach to identify and characterize these shocks directly. This approach, as outlined above, basically involves decomposing a given variable into two main components - permanent and transitory components - and analysing the relative effectiveness of each of these components in explaining fluctuations in the said variable. The common trends econometric technique as utilized here identifies and analyzes the long run effects of country-specific and global productivity shocks on fluctuations in investment and the current account. The theoretical framework utilized provides long run restrictions relevant for distinguishing between the global and country-specific productivity shocks. Our empirical results indicate that the common trends approach yields results that are comparable to those of Glick and Rogoff(1995) and goes beyond instantaneous least squares estimates by providing us with the possibility of dynamic analyses - which in this context is implemented using estimated impulse response functions of the effects of innovations in the identified productivity shocks; and forecast error decompositions - that are very crucial in empirical investigations of capital mobility and current account fluctuations. It turns out that the estimated productivity shocks are highly persistent and explain a significant proportion of variations in our variables of interest at the long run horizon.

1.2.2 Policy Shocks in the Open Economy

The chapters of this part of the dissertation dwell mainly on the effects of monetary and fiscal policy shocks on movements in exchange rates and the current account. Chapter five utilizes structural VARs in an attempt to address the *forward discount bias puzzle* and the *exchange rate puzzle* encountered in the literature. Uncovered interest parity requires that a fall in the domestic short-term interest rate (relative to the foreign short-term interest rate) which instigates an initial impact depreciation of the domestic currency should be followed by an appreciation since the reduced interest rate differential should be associated with an anticipation of domestic currency appreciation for agents to continue to hold domestic assets. Empirical findings not consistent with this requirement are said to yield a puzzle - the *forward discount bias puzzle*. This puzzle is encountered in Eichenbaum and Evans(1995) and Sims(1992) where (for some of the countries considered) positive domestic interest rate innovations are followed by large and persistent appreciations of the domestic currency. Secondly, could the *exchange rate puzzle* (which is the tendency of the domestic currency (of the non-US G-7 countries) to depreciate against the US dollar following contractionary domestic monetary policy shocks) that shows up in most empirical results in the literature be due to the specific schemes used to identify monetary policy shocks? This puzzle shows up in Grilli and Roubini(1995) as well as in Sims(1992) among others, where - in a non-structural VAR approach - innovations in short term interest rates and monetary aggregates are used in the respective papers as monetary policy shocks. We suggest that the puzzles could be due to the mode of identification of monetary policy shocks. To verify this claim the chapter benefits from the current state of the debate in the monetary policy literature on the identification of monetary policy shocks, taking the debate a step further by incorporating international policy interdependence into monetary policy identification schemes in an empirical analysis of the effects of monetary policy shocks on bilateral exchange rates.

The debate on the identification of monetary policy shocks is essentially about the appropriate empirical measure of the stance of monetary policy. As documented, among others, by Leeper and Gordon(1992) the use of monetary aggregates (the monetary base, M_1 and/or M_2 for instance) is plagued by the *liquidity puzzle* (where positive innovations in these aggregates are associated with interest rate increases, contrary to what theory suggests). To circumvent this puzzle, Bernanke and Blinder(1992) as well as Sims(1992) and many other researchers in the literature identify monetary policy shocks directly with innovations in short-term interest rates. However, Strongin(1995) argues that "*without any demonstrated empirical linkage between Federal Reserve actions and interest rate movements, it is unclear how innovations in interest rates can be reasonably attributed to monetary policy*" (p. 464, italics mine). It is this recognition that prompted Christiano et al(1992) to suggest innovations in nonborrowed reserves as monetary policy shocks. But this implies, contrary to the facts, that the Federal Reserve does accommodate neither reserve demand nor borrowing shocks. The Strongin(1995) measure of innovations in the mix of borrowed and nonborrowed reserves as monetary policy shocks is therefore intended to address this conceptual shortcoming of Christiano et al(1992). Strongin's findings are not conclusive either. Bernanke and Mihov(1995) suggest a linear combination of innovations in total reserves, nonborrowed reserves and the federal funds rate as monetary policy shocks. Their specification implies that the Federal Reserve accommodates both borrowing and demand shocks in its monetary policy measures. Their identification scheme encompasses those of Bernanke and Blinder(1992), Christiano et al(1992) and Strongin(1995). The structural VAR identification schemes (one based on Bernanke and Mihov's semi-structural scheme and the other, a structural identification scheme, very related to the former but essentially different in the sense that it introduces international monetary policy interdependence as a testable hypothesis) used in this chapter yield very plausible contemporaneous and dynamic estimates of the effects of monetary policy shocks on bilateral exchange rates for all the countries considered; and the puzzles largely disappear.

Chapter 6 of this dissertation integrates the effects of fiscal expansion on movements in exchange rates and the current account using a parsimonious combination of cointegration and vector-autoregressions. During the 1980s the long-term interest rates in the US rose, the dollar appreciated and the current account balance deteriorated following a steady increase in the US budget deficit that began in 1981. In Europe, since the unification of Germany, the movement of the value of the Deutsche mark vis-a-vis the US dollar and of long-term interest rates have been upwards due to the effects of the unification on the bond and foreign exchange markets. Branson(1993) argues that the appreciation of the deutsche mark - at that time - is mainly due to fiscal expansion by the federal government following the re-unification. Also the recent devaluation and the floating of the Swedish kronor took place at a time of relatively rising interest rates, a substantial and rising budget deficits and worsening current account situation.

These events suggest that there is some relationship among budget deficits, bilateral exchange rates, interest rate differentials and the current account. Indeed, the importance of budget deficits has been recognised and incorporated into many theoretical models, especially, Portfolio Balance Models (PBM) models. (See Branson (1983), (1984), (1985), (1988) and (1993); and Branson and Marcheese(1988)). In spite of the apparent recognition in the literature of the significance of the effects of fiscal policy on interest rates, the exchange rate and the current account, not much empirical work exists that integrates the effects of budget deficits in understanding real (bilateral) exchange rate and current account movements. The main argument advanced by Branson(1988) in explaining the misalignment of the dollar in the 1980s focuses on the expansionary fiscal policy with its attendant federal budget deficit. He argues that the expansionary budget program of the 1981 instigated a crowding out of both domestic private demand and foreign demand for the products of the United States as a result of higher real interest rates an appreciating dollar and consequently a deteriorating current account. The dollar had to depreciate after foreigners had accumulated dollars. Other researchers believe however (see for instance Feldstein(1986)), that the rise of the dollar in the 1980s was primarily due to the shift to a less inflationary monetary strategy, growing confidence in the US as a political "safe-haven" for investments following the election of Ronald Reagan as president, and/or a general perception by foreign investors of more favourable long term prospects for the US economy following the increased profitability of US corporations in the 1980s. In this chapter - chapter 6 - we attempt to add to the empirical literature on the effects of the fiscal policy innovations on exchange rate and current account fluctuations by incorporating budget deficits into the analysis of comovements in the macroeconomic variables of interest - budget deficits, money, interest rates, exchange rates and the current account balance. Our empirical results indicate that a larger proportion of short-term variations in the interest rate differential between financial centres as well as of variations in the current account balances of countries are accounted for by monetary innovations rather than by fiscal innovations. However, over the long run horizon, fluctuations in exchange rates are better explained by orthogonalized innovations in the fiscal policy variable.

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Chapter 2:

An Overview of Asset Market Based Theories of Exchange Rate Determination

2.1 Introduction

Most industrialised countries floated their currencies at the dawn of the breakdown of the Bretton Woods system of fixed exchange rates. This event offered researchers the opportunity to discuss the relative merits of fixed and freely floating exchange rate systems. Needless to say that at the root of all this is the development of theories explaining the observed large fluctuations in the current accounts of countries and the bilateral exchange rates between countries. At the core of these theories is the general class of asset market models[4]. These models are so labelled because they have certain characteristics reminiscent of asset markets - not only do they see exchange rates as relative prices of financial assets but also they give expectations a significant role in exchange rate determination. These expectations arise out of uncertainty as to future movements in exchange rates between currencies and the risk associated not only with the state of the economy but also with policy uncertainty in the economy in which the investment is carried out. This category of models encompasses the flexible price and sticky price monetary models as well as portfolio balance models the main elements of which are discussed in the subsections below. But before we discuss these models let us take a look at working definitions of some very commonly used concepts in exchange rate economics - Purchasing Power Parity, Covered Interest Parity and Uncovered Interest Parity.

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[4] Some researchers distinguish between equilibrium and disequilibrium models of exchange rate determination classifying sticky price models under the former and portfolio balance models under the latter category. Recent approaches using general equilibrium modelling (and purporting to explain the observed real exchange rate volatility solely by real productivity shocks) in explaining exchange rate volatility - see Stockman(1987), Mendoza(1991) and/or Backus *et al* (1992) - are classified under equilibrium models.

2.2 Concepts and Preliminary Definitions

2.2.1 Purchasing Power Parity (PPP)

Purchasing Power Parity (PPP) and Uncovered Interest Parity (UIP) are two of the most commonly used concepts in international finance. These concepts are important firstly because they underlie the bulk of models out there aiming at explaining the observed behaviour of exchange rates following the demise of Bretton Woods. Writing about the importance of the UIP condition McCallum states "... the main fact to be kept in mind is that it appears as a key behavioural relationship in virtually all of the prominent current-day models of exchange rate determination" - (McCallum(1994) p. 109). Secondly, empirical findings over the years have given birth to lots of discussions centred around these concepts. This is precisely because the empirical findings do not seem to be consistent with the theoretical predictions. Thus we have in the literature the Purchasing Power Parity Puzzle and the failure of the forward exchange rate to act as an unbiased predictor of the spot rate. These two puzzles have implications for our understanding of exchange rate fluctuations in response to changes in the underlying economic fundamentals and economic policy.

The thrust of Purchasing Power Parity is the assertion that in the absence of transportation costs, tariffs and other trade barriers the price of the same good in different countries should be the same if converted into a common currency. For instance, given this law of one price, we expect that if the price of one kilogram of coffee in the United States is 10 dollars then in Germany the price should be 15 deutsche marks if the exchange rate is 1.5 deutsche marks to one US dollar. Formally, for any good i we do expect the following relationship to hold

$$P_i = S_t P_i^* \qquad (2.1)$$

where P_i denotes the domestic price of good i, P_i* is the foreign currency price of the same good and finally S_t is the nominal exchange rate expressed as the domestic currency price of a unit of the foreign currency. The implication of this concept is then that the real exchange rate is constant and the price differential between countries are reflected in nominal exchange rate.

However the empirical evidence tends to indicate that relative nominal prices are less volatile that the nominal exchange rates and also large and persistent deviations from the law of one price are observed. Hence the puzzle. Figure 2.1 on page 24 shows the plots of the logarithms of the ratio of the relative consumer price index (CPI) and the nominal exchange rate for Canada and Germany relative to the US.

As we infer from figure 2.1 there is a significant difference between the variance of the exchange rate and that of the relative price indices for all the countries. This indicates a failure of Purchasing Power Parity. This failure has been explained by various researchers. Some researchers attribute this puzzling findings to transportation costs, tariffs and non-tariff trade barriers that drive a wedge between domestic and foreign price; preventing international arbitrage in the goods market to lead to equalization of these prices. Other researchers blame it on the short run stickiness of nominal prices and wages. Their view is as follows: as monetary shocks strike the nominal exchange rate the real exchange rate also changes in the short run. Thus preventing the expected constancy of the real exchange rate as predicted by Purchasing Power Parity. But if this is true then one would expect a convergence to PPP once prices and nominal wages begin to adjust to the shock after the period of inertia. This is however not observed in most empirical work in the literature. This puzzling empirical evidence raises lots of questions as to the nature of shocks driving real exchange rate movements. Following Dornbusch(1976) most researchers attribute a large role to financial and monetary innovations. Shocks to productivity and preferences, as the conventional thinking goes, are not volatile enough to explain the observed short term volatility of exchange rates. Clarida and Gali(1994) in an attempt to verify this investigate the role of nominal and real shocks in explaining the observed short term real exchange rate movements. Their findings

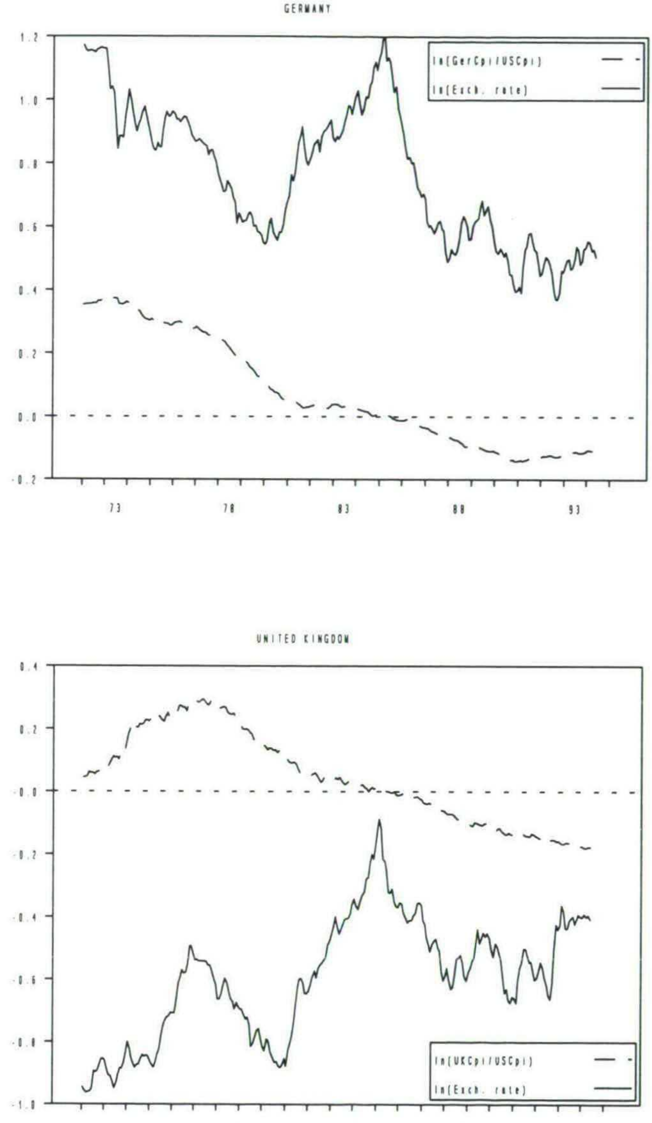

Figure 2.1: Exchange Rates and Relative (CPI) Price Ratios. Evidence of Short-Term PPP Failure.

indicate that about 45% and 34% of the forecast error variance of the deutsche mark/dollar exchange rate and of the yen/dollar real exchange rates respectively are explained by monetary shocks. Rogers(1995) finds that real shocks explain roughly 50% of the forecast error variance of the british pound/ US dollar real rate. Thus the puzzle is far from resolved - it is still not very clear what the relative contributions of real and nominal shocks are in explaining the observed short run real exchange rate variability contrary to the constancy predicted by PPP.

2.2.2 Uncovered Interest Parity (UIP)

The forward exchange market is an institution for trading in foreign exchange for delivery in the future. The rate of exchange at which a contract is signed to purchase or sell foreign currency in the future is called the forward exchange rate. This forward rate is made up of two parts - the current spot exchange rate plus (a discount) or minus (a premium) a margin. The forward rate is determined in a free market by the interest rate differential between the financial centres involved. Suppose the domestic (say Germany) interest rate is 3% per annum and that of a foreign country (say the US) is 5% per annum. Clearly then German investors could increase their interest earning by buying dollars and depositing the amount in the US for a year. But then after a year the deutsche mark may have depreciated by more that the interest rate differential between the US and Germany. In other words the German investors would make a loss when compared with the yield on a German deposit over a year. To overcome this risk the investors could undertake two transactions simultaneously. They could buy US dollars spot, invest the amount in the US and at the same time sign a contract to sell the yield for domestic currency forward - that is at a forward rate. So long as the yield on the investment is greater than that on a comparable investment in Germany the investors could continue to make this risk-free pair of transactions and earn a guaranteed profit. This arbitrage would continue to be earned until the yield on similar assets/investments are equalized between Germany and the US - that is until

$$(1 + r_t) = (1 + r_t^*)\frac{f_t}{S_t} \tag{2.2}$$

where r_t is the domestic (German) interest rate, r^*_t is the US interest rate, S_t is the spot exchange rate (i.e. deutsche marks per US dollar) and f_t denotes the forward exchange rate. Equation (2.2) above is the covered interest parity condition. Uncovered Interest Parity (UIP) holds when the forward exchange market is efficient and the forward rate is an unbiased predictor of the future spot rate. The forward market is said to be efficient when prices in the market reflect all information available to all agents in the market. It should be impossible as well to make risk-free profits by knowing better than the market.

Forward exchange market efficiency thus implies the joint proposition that participants in the market are endowed with rational expectations and are risk-neutral. The former proposition - of rational expectations - can be expressed formally as

$$\Delta S_{t+k} = \Delta S^e_{t+k} + \varepsilon_{t+k}, \qquad \textit{where} \qquad \Delta S^e_{t+k} = E[\Delta S_{t+k} | I_t] \tag{2.3}$$

where S_t is the logarithm of the spot exchange rate defined as the domestic currency price of foreign currency, S^e_{t+k} is the expected value of the spot rate k periods ahead and I_t is the information set on which expectations are based. E is the expectations operator whereas ε_{t+k} is a random forecast error that is orthogonal to the agents' information set. If agents are risk-neutral one would expect the forward exchange rate to be driven to equality with the expected value of the k-periods ahead spot rate - that is $f_t = S^e_{t+k}$. Risk-aversion drives a wedge between these two and the proposition of rational expectations and risk-aversion could be expressed by the help of equation (2.3) as

$$f_t - s_t = \Delta s^e_{t+k} + \rho_t = \Delta s_{t+k} + \epsilon_{t+k} \qquad (2.4)$$

where $\varepsilon_{t+k} = \rho_t - \vartheta_{t+k}$, ρ_t is a time-varying risk premium term and ϑ_{t+k} is a random forecast error term. The expression $f_t - s_t$ is the forward premium.

Empirical tests of uncovered interest parity (UIP) could be formulated as a test of the ability of the forward rate to act as an unbiased predictor of the future spot rate. Alternatively one could test the hypothesis of unbiasedness of the forward premium in predicting changes in the spot rate. The two hypotheses could be expressed as a test of the null hypothesis of $\beta = 1$ in the respective formulations $S_t = \alpha + \beta f_{t-1} + \varepsilon_t$ *and* $S_t - S_{t-1} = \alpha + \beta(f_{t-1} - S_{t-1}) + \varepsilon_t$. It turns out that researchers over the years have revealed sharp contrasts in the estimates of the 'unbiasedness' coefficient β depending on the formulation adopted. See for instance MacDonald(1988), McCallum(1994) and the survey data results of Frankel and Froot(1990) for a survey of the empirical regularities obtained so far in the literature. Whereas in specifications of the former type estimates of approximately 1 are obtained for the biasedness coefficient - supporting unbiasedness of the forward rate - empirical tests adopting the second formulation get estimates in the range of -3 and -4 refuting the hypothesis that the forward premium is an unbiased predictor of future spot rate changes.

2.3 The Flexible Price Monetary Model

The flexible-price model assumes the existence of tradable goods and that equilibrium in the tradable goods market ensures that in the absence of transportation costs and other distortions, when there exist no incentives for further trade flows, prices denoted in a common currency are equalized and purchasing power parity (PPP) holds - at least to a disturbance term - and

$$s_t = p_t - p_t^* + v_t \tag{2.5}$$

where s_t is the logarithm of the nominal exchange rate - i.e. the domestic price of foreign currency, p_t denotes the logarithm of the domestic price level, p_t^* denotes the logarithm of the foreign price level and v_t is assumed stationary disturbance term capturing short run deviations from PPP. The other major relations of the flexible-price model include the usual money demand functions of the domestic and foreign country as presented below (all foreign country coefficients and variables are denoted with asterisks)

$$m_t = p_t + \phi y_t - \lambda r_t + \xi_t \tag{2.6}$$

$$m_t^* = p_t^* + \phi^* y_t^* - \lambda^* r_t^* + \xi_t^* \tag{2.7}$$

where m_t, p_t and y_t denote respectively the logarithm of the domestic money stock, price level and income whereas ϕ and λ are positive coefficients. Given equations (2.5) - (2.7) we can solve for the nominal exchange rate as a function of the relative money supplies, income and the interest rate once we assume equality between the domestic and foreign coefficients - i.e. for $\phi = \phi^*$ and $\lambda = \lambda^*$ - as represented below

$$s_t = (m_t - m_t^*) - \phi(y_t - y_t^*) + \lambda(r_t - r_t^*) - (\xi_t - \xi_t^*) + v_t \tag{2.8}$$

One crucial assumption of the flexible-price model is that of perfect substitutability of domestic and foreign assets - or equivalently the assumption that uncovered interest parity (UIP) holds such that domestic interest rate reductions should be associated with an anticipated appreciation of the domestic currency in order to give incentives for continued holding of domestic currency denominated assets. This condition is usually represented by the relationship

$$r_t - r_t^* = E[\Delta s_{t+1}|I_t] \tag{2.9}$$

where E is the expectations operator and I_t denotes the information set of agents up to and including the present period. Substituting this relation into (2.8) yields the fundamentals-equation

$$s_t = FU_t + \lambda E[\Delta s_{t+1}|I_t] + \xi_t^1 \tag{2.10}$$

where FU_t denotes a combination of variables $(m_t - m_t^*)$ and $-\phi(y_t - y_t^*)$, known as the *fundamentals*, and ξ^1_t is a combination of stationary disturbances. Thus equation (2.10) above conveys the asset-market idea of modelling the nominal exchange rate as a function of future expectations of the fundamentals (including among others the relative foreign and domestic asset supplies).

If we ignore the short run stationary disturbances and hold foreign variables constant for convenience we observe from equation (2.8) that increases in the domestic money stock instigate domestic price level increases that necessitate the depreciation of the domestic currency in conformity with the purchasing power parity relation. Although domestic interest rate increases make domestic assets more attractive to domestic and foreign investors alike, the reduced demand for domestic currency (following the shift to domestic financial assets other than money) instigates its depreciation. Increases in the domestic growth rate of income encourage domestic agents to increase their demand for real money balances leading to domestic price falls (as demand pressure in the goods market is reduced) until market equilibrium is attained. As the domestic price level falls, purchasing power parity ensures the appreciation of the domestic currency.

This approach to exchange rate modelling has been criticised not only from theoretical standpoints but also from empirical failures of aspects of its crucial elements. On the theoretical front it is contended that open economy macroeconomics is essentially about the behaviour of agents in additional markets - the goods, labour, foreign exchange and the domestic and foreign bonds markets - other than just the money market effecting capital flows and affecting exchange rate and current account fluctuations. Empirical tests of the crucial assumption of PPP as well as the theoretical predictions of the flexible-price model yield unfavourable results. For an excellent review of the empirical performance of the flexible-price model see for instance Taylor(1995). In response to these criticisms - especially with respect to the contention that PPP does not hold (at least not in the short run) - naturally emerges the sticky-price literature championed by Dornbusch(1976). This is our subject of study in the next subsection.

2.4 The Sticky-Price Monetary Model

The main feature of this approach is linked with the proportionality between the exchange rate and the contemporaneous interest rate differential. Under the assumption of certainty equivalence, the uncovered interest parity relation (UIP) can be written in continuous time form as

$$\dot{s}_t = r_t - r_t^* \tag{2.11}$$

We then substitute a Phillips curve equation (2.12) below where the flexible-price model assumes continuous PPP.

$$\dot{p}_t = \pi(d_t - y_t) \qquad (2.12)$$

where d_t denotes aggregate demand and is dependent on real income, y_t, the real exchange rate, $s_t - p_t + p^*_t$, the nominal interest rate, r_t, and a fiscal shock term, g_t, as presented below:

$$d_t = \gamma y_t + \delta(s_t - p_t) - \sigma r_t + g_t \qquad (2.13)$$

where we assume the foreign price level constant. Given that the money market clears and asset yields are equalised such that equation (2.11) holds we obtain

$$p_t - m_t = -\theta y_t + \lambda r_t^* + \dot{s}_t \qquad (2.14)$$

The long run equation for the price level can be obtained by setting $\dot{s}_t = 0$ into equation (2.14). The resulting expression can be substituted back into equation (2.14) yielding the short run relation between the nominal exchange rate and the price level as below:

$$\dot{s}_t = \frac{1}{\lambda}(p_t - \bar{p}) \qquad (2.15)$$

which is one of the important relations of the sticky price monetary model. To derive the second relevant equation note that from equations (2.14) and (2.15) above we can easily obtain an expression for the change in the price level, $\dot{p}_t$. In the long run (i.e when the prices and the exchange rate are constant and interest rates are equalized) we obtain, after substituting equations (2.13) and (2.14) into equation (2.12), a relationship for the long run exchange rate as a function of real output, the foreign interest rate, the fiscal shock and the long run price level as

$$\bar{s} = \frac{(1 - \gamma)}{\delta}y + \frac{\sigma}{\delta}r^* - \frac{i}{\delta}g + \bar{p} \qquad (2.16)$$

where the long run price level is as defined in equation (2.15) above. It is easy to show further that the change in the price level can be expressed as a function of the deviations of the nominal exchange rate and the price level from their respective long run levels as in equation (2.17) below:

$$\dot{p}_t = \pi\delta(s_t - \bar{s}) + \pi(\delta + \frac{\sigma}{\lambda})(p_t - \bar{p}) \qquad (2.17)$$

The two equations - (2.15) and (2.17) - define the dynamics of the sticky price monetary model. Since the determinant of this system of equations - $\begin{vmatrix} 0 & \frac{1}{\lambda} \\ \pi\delta & \pi(\delta + \frac{\sigma}{\lambda}) \end{vmatrix} = -\frac{\pi\delta}{\lambda}$ - is negative, the system has a unique convergent saddle path as shown in Figure 2.2 below where SS is the derived negatively sloped saddle path.

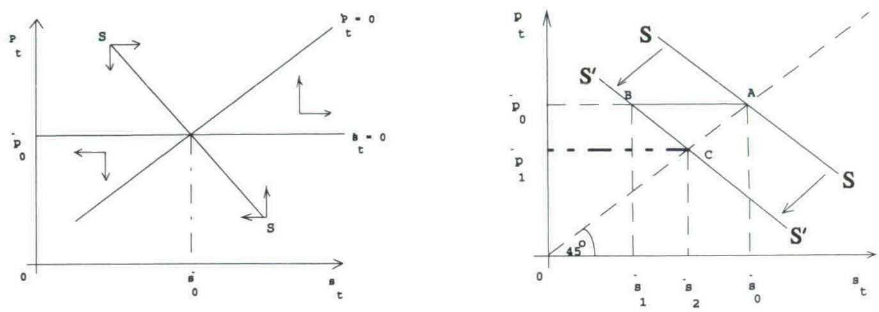

Figure 2.2: Exchange Rate Overshooting in the Sticky-Price Monetary Model

To help us discuss the effects of monetary policy within this framework we use figure 2.2 above to show the change in the saddle path (SS is the original saddle path and S'S' the new saddle path) following the policy shock. Consider the effects of a contractionary monetary policy. The long run exchange rate as given by equation (2.12) is dependent on the long run level of domestic prices, real income and fiscal shocks. The monetary contraction, if expected to persist, will put both the goods and the asset markets into disequilibrium. Since PPP holds as a long run relationship and foreign prices are held constant the long run exchange rate will appreciate proportionately along the 45 degree line until a new equilibrium is attained at point C where the price level has decreased and the domestic currency appreciated. However, because prices are sluggish, price adjustment lags behind the exchange rate adjustment. This implies that, given long run PPP and constant foreign prices, the exchange rate jumps down to point B on the new saddle path following contractionary monetary policy and prices adjust slowly bringing the exchange rate up with it (i.e depreciating the domestic currency) until point C is reached where equilibrium is re-installed in both the goods and the assets markets. The overall effect of the contractionary monetary policy is therefore a long run appreciation, $\bar{s}_0 - \bar{s}_1$, with an initial overshooting, $\bar{s}_2 - \bar{s}_1$. This overshooting phenomenon is explained by the fact that the positive domestic monetary shock drives a wedge between returns on domestic and foreign assets - in this case making domestic assets less attractive. In order to keep both categories of assets equally attractive the price of foreign exchange must increase sufficiently to generate an expectation of a decline in the future - reducing the returns on foreign assets and bringing them again back in equality with those on domestic assets.

2.5 Portfolio Balance Models

The Portfolio Balance model introduces the bond and the foreign exchange markets into the analysis of the determinants of the exchange rate and the current account and allows imbalances to have a feedback effect on the long run equilibrium through wealth. Thus the model allows for full interaction among the exchange rate, the current account balance and the level of wealth. The key modification introduced by a number of authors in this area is the assumption that domestic and foreign financial assets are not perfect substitutes. See Branson(1984), Branson and Hendersen(1985), and Kouri(1976) for instance for a comprehensive treatment of the portfolio balance model (PBM). This assumption motivates the inclusion of a risk premium in the uncovered interest parity relation as well as the introduction of the supply of bonds and other non-monetary assets into the analysis of the determination of, and fluctuations in, the exchange rate and the current account balance.

We discuss the simplest rational expectations form of the model as represented in the system of equations (2.18) - (2.22) below. The net financial wealth of the private sector (W_t) is divided into three components: money (M_t), domestically issued bonds (B_t) and foreign bonds denominated in foreign currency (F_t). The current account balance by definition depicts the rate of accumulation of foreign assets over time. Foreign and domestic interest rates, as before, are denoted r_t and r^*_t respectively, and $\dot{s}_t$ is the expected depreciation of the domestic currency.

$$W_t = M_t + B_t + s_t F_t \,, \qquad (2.18)$$

$$M_t = m(r_t, r_t^* + \dot{s}_t) W_t \qquad m_r < 0, \quad m_{r_t^* + \dot{s}_t} < 0, \qquad (2.19)$$

$$B_t = b(r_t, r_t^* + \dot{s}_t) W_t \qquad b_r > 0, \quad b_{r_t^* + \dot{s}_t} < 0, \qquad (2.20)$$

$$s_t F_t = f(r_t, r_t^* + \dot{s}_t) W_t \qquad f_r < 0, \quad f_{r_t^* + \dot{s}_t} > 0, \qquad (2.21)$$

$$\dot{F}_t = n(\frac{s_t}{p_t}, z_t) + r_t^* F_t \qquad n_{\frac{s_t}{p_t}} > 0, \quad n_z > 0. \qquad (2.22)$$

Note that the variables in the equations above are in levels and hence S_t is the domestic price of foreign currency. Also, as generally specified, $X_w = \delta X / \delta w_t$. The solution to the rational expectations specification yields among others the result that unanticipated positive real disturbances that improve the current account balance create expectations of an appreciation of the domestic currency to bring the current account back to equilibrium - see equation (2.22). An unanticipated monetary policy shock causes a jump in the exchange rate followed by a gradual adjustment into a new long run equilibrium with a depreciated domestic currency. To formally analyze these effects we divide equations (2.19) and (2.21) by W_t and totally differentiate them, holding r^*_t constant. This yields the following equations

$$d(\frac{M}{W}) = m_r dr + m_s ds$$
$$d(\frac{sf}{W}) = f_r dr + f_s ds \qquad (2.23)$$

which can then be solved using matrix algebra and yielding explicit expressions for *dr* and *ds* as below:

$$\begin{bmatrix} dr \\ ds \end{bmatrix} = \frac{1}{(f_r m_s - m_r f_s)} \begin{bmatrix} m_s & -f_s \\ -m_r & f_r \end{bmatrix} \begin{bmatrix} d(\frac{sf}{W}) \\ d(\frac{M}{W}) \end{bmatrix} \qquad (2.24)$$

From the solution for ds we can express the expected change in the exchange rate as a function of relative shares of foreign and domestic financial assets in the agents' portfolio - thus

$$\dot{s} = \phi(\frac{sF}{W}, \frac{M}{W}), \quad \textit{where} \quad \phi_1 > 0 \textit{ and } \phi_2 < 0. \qquad (2.25)$$

We solve the dynamic equations (2.22) and (2.25) for the equilibrium levels of *s* and *F*. We totally differentiate these equations to obtain the slopes of the respective loci as follows:

$$d(\frac{s}{dF})\Big|_{\dot{s}=0} = -\frac{s}{F} \quad \textit{and} \quad d(\frac{s}{dF})\Big|_{\dot{s}=0} = -\frac{r_t^*}{n_{\frac{s}{p}}} \qquad (2.26)$$

Given these slopes and the behaviour of the exchange rate and the foreign asset holdings off their respective loci, saddle point equilibrium requires that $-s/F$ be less than $-r^*/n_{s/p}$; which is but the Marshall-Lerner condition requiring that $sn_{s/p}/n$ be greater than one. The equilibrium paths for s and F (given that the Marshall-Lerner condition holds) are depicted in figure 2.3 below.

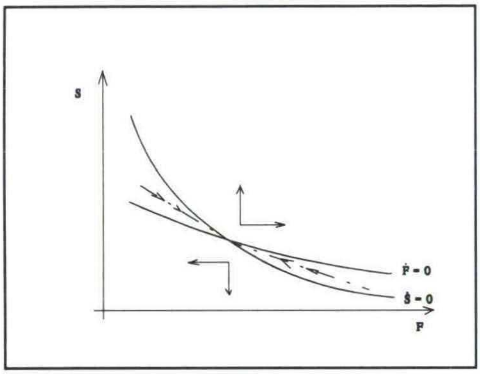

Figure 2.3a: Equilibrium path for S_t and F_t

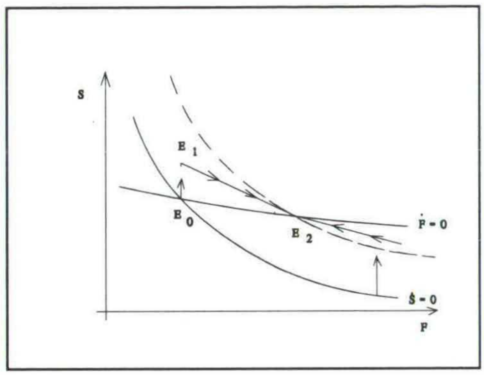

Figure 2.3b: Dynamic effects of an unanticipated monetary policy shock

Now let us consider the dynamic effects of an unanticipated expansionary monetary policy shock. The rise in M shifts the expected depreciation locus upwards since an increase in M requires an offsetting exchange rate depreciation or an increase in foreign asset holdings to hold $\dot{s}_t = 0$. The new long run equilibrium is at E_2. However given that the shock is unanticipated agents jump initially to E_1 on the new saddle path associated with the new locus and move gradually towards the new equilibrium at E_2. Thus, the adjustment process after the unanticipated monetary expansion clearly involves the exchange rate overshooting its long run equilibrium level. Real disturbances such as technological shocks and terms of trade improvements (considered in the context of this model as effects of a permanent shift in z_t) do shift the $\dot{F}_t = 0$ locus downwards leading to exchange rate undershooting during the movement to the long run equilibrium that involves an appreciation of the domestic currency.

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Chapter 3

Stochastic Trends and Fluctuations in the Interest Rate, Exchange Rate and the Current Account Balance: An Empirical Investigation.[1]

3.1 Introduction

The period after the breakdown of the Bretton Woods agreement in the early 1970s has been marked by generalized floating of the world's major currencies against each other. Also, since the early 1980s the world economy has been plagued by large and uncoordinated innovations in monetary and fiscal variables as well as by high and volatile real interest rates, large movements in currency exchange rates and significant variations in current account balances. The post Bretton Woods period of generalized floating exchange rates among the major currencies of the world provides the opportunity for observation and analysis of the multivariate relationship between real income, an index of expansionary fiscal policy (say the budgetary deficit), money supply, the interest rate differential among financial centers, real exchange rates and the current account balance. In response to this opportunity a substantial and wide-range literature on the theoretical and the empirical linkages among the variables mentioned above has emerged.

One strand of the literature consists of two classes of asset market models. The first class consists of models that fall under the Monetary Approach to exchange rate modelling which encompasses sticky and flexible price models characterized by the implicit assumption of perfect substitutability between domestic and foreign money and bonds. The second class of models - the Portfolio Balance Model (PBM) does not make the assumption of perfect

[1] This chapter is a slightly revised version of Kumah and Ibrahim(1996) published in ***Economic Modelling***, 13(3), 1996 pages 383-406. We are very grateful to William H. Branson C-H Siven, Goran Eriksson and an anonymous referee constructive comments. However any remaining unclarities and/or errors remain our responsibility.

substitutability between domestic and foreign financial assets. The PBM emphasizes wealth effects on asset demands and the role of the exchange rate and expectations about its future movements in the asset demand decision. Some versions of the PBM introduce the current account balance in its role as allocating wealth among countries. A surplus/deficit in the current account represents a rise/fall in net domestic holdings of foreign assets which in turn affects the level of wealth and the real exchange rate. Extending the PBM to include rational expectations Branson(1983 and 1984) demonstrates that real shocks generate monotonic adjustment of the exchange rate while monetary shocks generate some degree of exchange rate overshooting.

The other strand of the literature on the effects of real and nominal shocks on the real exchange rate and the current account balance comes under the heading dynamic intertemporal approach to the current account balance. This approach using the assumption of a forward looking utility maximizing representative agent, emphasizes the effects of real factors (such as productivity), the terms of trade, taxation and government spending on the current account balance - through their intertemporal substitution effects on consumption, production and investment.

The empirical results from this class of asset market models of exchange rate determination has not been encouraging so far. Empirical tests of the implications of the intertemporal approach is not widespread either. This may be due to the intractable nature of some of the models and the high demand placed on data. Razin and Rose(1992), and Mendoza(1991) obtain some encouraging results in their attempts to test some of the key hypothesis/implications of the model. Due to the non-conclusive nature of the empirical results from these classes of models we adopt a multivariate data analysis approach - a Common Trends Approach - which allows the data to speak for themselves. This is an approach closely related to international business cycle models of Stockman(1983), Mussa (1986), Baxter and Stockman(1989), Backus, Kehoe, and Kydland(1992) and Meltzer(1986 and 1992). Recent work by Eichenbaum and Evans(1995) on the effects of monetary policy shocks on the nominal and real exchange rate is in the same fashion. One general characteristic of these works is that they adopt a simple vectorautoregression (VAR) approach even though the authors are aware that the exchange rate as well as the current account balance are driven by real as well as nominal domestic and foreign trends. This chapter is an attempt at incorporating these trends into the analysis of fluctuations in the exchange rate and the current account balance in response to policy shocks. An early attempt to illustrate the use of this approach in an open economy setting was by Warne(1990) - this chapter closely follows his approach and uses the common trends estimation method developed and programmed by the same author.

The rest of the chapter is organized as follows. The next section introduces the common trends literature and shows its link with cointegration. The third section performs the preliminary data analysis and estimation of the common trends model using a six variable vector. Our empirical findings are summarized in the final subsection of this section. The final section provides a summary our empirical findings and concludes the chapter.

3.2 Common Trends in the Open Economy

In this section, we present a common trends model to characterize fluctuations in some selected key variables (including, among others, interest rate differentials among countries, the exchange rate, and the current account balance) in the open economy. According to the common trends literature[2] any n x 1 dimensional macroeconomic time series X_t can be decomposed into permanent (X_t^p) and transitory (X_t^s) components i.e.

[2] See for instance Blanchard and Quah (1989), King, Plosser, Stock and Watson(1991), and Stock and Watson (1988).

$$X_t = X_t^p + X_t^s \tag{3.1}$$

which is but a decomposition of the cointegrated[3] Wold vector moving average representation of the X_t series:

$$(1 - L)\, X_t = \delta + C(L)\varepsilon_t \tag{3.2}$$

where $C(L) = I_n + C_1 L + C_2 L^2 + \ldots$ (an invertible lag polynomial), ε_t is white-noise with a zero mean vector (i.e. $E[\varepsilon_t] = 0$) and a positive definite covariance matrix, $\Xi = E[\varepsilon_t \varepsilon_t']$, and L is the lag operator such $L^i \chi_t = \chi_{t-i}$.

Under fairly general conditions it is possible (using matrix algebra) to find another polynomial $C^*(L)$ such that

$$C(L) = C(1) + (1 - L)C*(L) \tag{3.3}$$

where the polynomial matrix $C^*(L)$ is given by

$$\begin{aligned} C^*(L) &= (1 - L)^{-1}(C(L) - C(1)) \\ &= \sum_{j=0}^{\infty}\left(\sum_{i=0}^{j}(C_i - C(1)\right)L^j = \sum_{j=0}^{\infty}\left(\sum_{i=j+1}^{\infty} C_i\right)L^j = -\sum_{j=0}^{\infty} C_j^* L^j . \end{aligned}$$

To derive the common trends representation of equation (3.2) we require certain results from the Granger Representation Theorem as stated below.

[3] We allow the X_t vector to be cointegrated of order (1,1) in an Engle-Granger sense (See Engle and Granger (1987). However, there is a slight difference in the usage of the terminology here since we have some $I(0)$ variables in X_t .

3.2.1 The Granger Representation Theorem (GRT)

Let (3.2) be cointegrated with $d = 1$ and $b = 1$ (i.e X_t is $CI(1,1)$) with cointegrating rank r[4]. Then

i) $C(1)$ is of rank $n - r$ where n is the number of variables in X_t and r is the number of cointegrating vectors.

ii) there exists an ARMA representation

$$\Pi(L)X_t = d(L)\varepsilon_t \qquad (3.4)$$

where $\Pi(0) = I_n$, rank of $\Pi(1)$ is r, and $d(L)$ is a scala lag polynomial.

iii) there exist $n \times r$ matrices β and α, of rank r such that

$$\beta' C(1) = 0,$$

$$C(1)\alpha = 0, \text{ and}$$

$$\Pi(1) = \alpha\beta'$$

iv) there exists an Error Correction Representation with $Z_t = \beta' X_t$, an $r \times 1$ vector of stationary variables:

$$\Pi^*(L)(1 - L)X_t = \delta + \alpha Z_{t-1} + d(L)\varepsilon_t \qquad (3.5)$$

with $\Pi^*(0) = I_n$, an $n \times n$ identity matrix.

Utilizing equation (3.3) and the usual convention of letting ε_s be zero (for $s \leq 0$) and X_0 representing the non-random initial value of X_t,, then by recursive substitution[5] of equation (3.2) we obtain,

$$X_t = X_0 + \delta t + C(1)(1 + L + L^2 + ... + L^{t-1})\varepsilon_t + C^*(L)\varepsilon_t \qquad (3.6)$$

---

[4] The components of the vector X_t are said to be cointegrated of order d, b, denoted $X_t \sim CI(d\text{-}b)$ if all the components of X_t are $I(d)$; and there exists a vector $\alpha \neq 0$ so that $Z_t = \alpha' X_t \sim CI(d\text{-}b)$, $b > 0$. The vector α is called the cointegrating vector.

[5] See Stock and Watson(1988) for the procedure applied here.

Since δ is one of the vectors in the nullspace of $C(1)$, it can be expressed as a linear combination of the columns of $C(1)$, say $\delta = C(1)\upsilon$ where υ is an $n \times 1$ vector. Substituting this expression into equation (1.3) above we obtain

$$X_t = X_0 + C(1)[t\upsilon + \sum_{i=0}^{t}\varepsilon_i] + C^*(L)\varepsilon_t$$

Denoting the expression in the square brackets (the random-walk component) by $\tau_t = \mu + \tau_{t-1} + \varphi_t$, we obtain, after some manipulations, the expression

$$\begin{aligned} X_t &= X_0 + A\tau_t + C^*(L)\varepsilon_t \\ &= X_0 + A[\mu + \tau_{t-1} + \varphi_t] + C^*(L)\varepsilon_t \end{aligned} \qquad (3.7)$$

where $C(1)\varepsilon_t = A\varphi_t$, $C(1)\delta = A\mu$, and hence $C(1)\Xi C(1)' = A\Phi A'$. Equation (3.7) is the multivariate version of the Beveridge-Nelson decomposition[6] of the cointegrated vector moving average representation, (3.1), expressing X_t as a linear combination of $n - r$ linearly independent stochastic trends (the common trends that have permanent effects of X_t) and transitory components, $C^*(L)\varepsilon_t$ which are stationary. X_0 contains the initial values of X_t.

$$\tau_t = \mu + \tau_{t-1} + \varphi_t \qquad (3.8)$$

is a $k \times 1$ vector of non-stationary variables (with structural white-noise shock vector, φ_t, with the variance-covariance matrix, Φ) that drive the system, A is a rank k matrix of coefficients to be characterized (by the aid of an underlying theory) and estimated.

The link between cointegration and common trends can be brought to the fore if we realize that if a vector of n variables, X_t - not all the members of which are $I(1)$ variables - is

[6] See Beveridge and Nelson (1981) or Stock and Watson (1988) for the details of this decomposition.

cointegrated of order (1,1) then there exists a common trends representation with $n - r$ linearly independent stochastic trends (or, equivalently, there exists $n - r$ non-stationary variables (unit roots), where n denotes the number of variables in X_t and r is the number of cointegrating vectors). In our search for the common trends therefore we make use of the tools of cointegration, especially the Johansen's method of testing for the cointegration rank.

3.3 Empirical Results

The data utilized for the analysis includes six vectors of quarterly seasonally unadjusted series for the post Bretton-Woods period 1975:1 to 1992:4 (obtained from various issues of International Financial Statistics) for Germany, Japan, Sweden, the United Kingdom and the United States of America. For each country, the variables used include the logarithm of real gross national product ($log\ y_t$), an index for expansionary fiscal policy ($bdef_t$)[7], the logarithm of the broad money supply ($log\ m3_t$), interest rate differential (measured as the difference between the domestic rate and the US federal funds rate ($i_t^D - i_t^{US}$), the nominal exchange rate (e_t), and the current account balance (ca_t). The expansionary fiscal policy index and the logarithm of money supply are denoted exogenous variables in the system. Hence the X_t vector used takes the form $[\ log\ y_t \quad (\ i_t^D - i_t^{US}\ ) \quad e_t \quad ca_t\ ]'$.

[7] For all countries, except Japan, the index used is the structural budgetary deficit. For Japan government consumption expenditures are used as a proxy for the structural budgetary deficit because of problems of data availability.

3.3.1 Empirical Tests for Common Stochastic Trends

In this section we review both the univariate and multivariate stochastic properties of the data. Common trends analysis presumes that some of the vectors of the system are non-stationary variables (i.e. there exists k unit-roots in the system). The univariate properties of the data are investigated using the Dickey-Fuller (DF) test, and the Stock-Watson (Q_f) test. The results of these tests are not presented here however due to complexities of the presentation which involves the statement of the levels of augmentation for each series in the data set for each country. We shall therefore rely only on the multivariate results derived from the Johansen's maximum likelihood procedure below.

Preliminary Cointegration Analysis and test for Common Trends

We now turn to the investigation of the multivariate stochastic properties of the data - more specifically, we perform stationarity tests using Johansen's method and from the results of the test, determine the number of common trends in the system. Let the n variables under investigation follow a VAR process of order p as below.

$$X_t = \Pi_1 X_{t-1} + \Pi_2 X_{t-2} + \dots + \Pi_p X_{t-p} + \mu + \Psi D_t + \epsilon_t \tag{3.9}$$

where the elements of the ε_t vector are i.i.d. normal with a zero mean and a constant variance. We re-parameterize equation (3.5) into the Error Correction representation which we specify here (without derivation[8]) as:

[8] See Juselius, K. and S. Johansen (1990) for the derivation of this Error Correction specification.

$$\Delta X_t = \sum_{1=1}^{p} \Gamma_i \Delta X_{t-i} + \Pi X_{t-p} + \mu + \Psi D_t + \epsilon_t \qquad (3.10)$$

where

$$\Gamma_i = -[I_n - \Pi_1 - \ldots - \Pi_i], \quad i = 1, \ldots, k-1.$$
$$\Pi = -[I_n - \Pi_1 - \ldots - \Pi_p] \quad \textit{and}$$

I_n is an $n \times n$ identity matrix. The hypothesis to be tested (conditional on the lag length, p) - which is equivalent to a test for the existence of k stochastic common trends - is that there exist reduced rank matrices α and β such that $\Pi(1) = \alpha\beta'$ where α and β are nxr matrices and $r = n - k$. To determine the lag length, p we use the Akaike Information Criteria (AIC), the Schwartz Information Criteria (SIC) and the Iterated Log Criteria (ILC). The AIC test picked a lag length of six whereas the other two tests picked one lag length for the unrestricted vector autoregressions. Because of this inconsistency in the results of these statistics we decided to let the multivariate as well as the univariate statistics[9] of the residuals of the Vector Error Correction Model determine the lag length for a given number of cointegrating vectors (in this case two). The results of the test for the number of cointegrating vectors are presented in table 3.1 below.

[9] The multivariate as well as the univariate statistics are not presented here but are available with the authors and can be obtained upon request.

Table 3.1: Likelihood Ratio Tests for the number of Cointegrating vectors

A.The Maximum Eigenvalue (λ_{max}) Criteria.

Hypothesis		Test Statistics			
H_0	H_1	Germany	Japan	Sweden	UK
r=0	r=1	34.763*	37.519*	33.262*	42.263*
r≤1	r=2	10.876	32.040*	18.667**	26.461*
r≤2	r=3	2.430	11.071	10.930	10.180
r≤3	r=4	0.101	3.761	2.357	4.076

B. The Trace Criteria.

Hypothesis		Test Statistics			
H_0	H_1	Germany	Japan	Sweden	UK
r=0	r≤1	48.169*	84.391*	65.216*	82.981*
r≤1	r≤2	13.406	46.872*	31.954*	40.718*
r≤2	r≤3	2.530	14.831	13.287	14.256
r≤3	r≤4	0.101	3.761	2.357	4.076

Note: *The estimated model had six lags, an intercept and three seasonal dummy variables. Hence Table 1 of Osterwald-Lenum (1990) page 468 was used. The asterisk indicate the following levels of significance: * 5% and ** 10%.*

In the exception of Germany (where the results above seem to suggest the existence of only one cointegrating vector) the results presented in the table above seem to accept the existence of two cointegrating vectors for all countries considered at the 5% significance level. The empirical results for Germany notwithstanding we utilize two cointegrating vectors in all subsequent analyses for all countries considered. This implies that we have two common stochastic trends in the data set. These two stochastic trends are identified respectively as real domestic trend (or domestic technological/supply shocks), τ_R, and domestic nominal trends, τ_N. In order to be able to estimate the A matrix in equation (3.7) we need to characterize it further. In the common trends tradition, if X_t has k common trends then the A matrix can be written as the product of two matrices, A_0 and π, i.e

$$A = A_0\pi \tag{3.11}$$

where A_0 is an $n \times k$ long run impact matrix with known parameters (characterizing the effects of each of the stochastic trends on the respective elements of the X_t vector) chosen such that $\alpha' A_0 = 0$, and π is a $k \times k$ lower triangular matrix of unknown parameters. For each country we specify the A_0 matrix as

$$A_0 = \begin{bmatrix} 1 & 0 \\ 1 & 1 \\ a_{31} & a_{32} \\ a_{41} & a_{42} \end{bmatrix} \tag{3.12}$$

This matrix (incorporating the required $k\mathrm{x}(k-1)/2$ restrictions for the identification of the common trends model) ensures that, in the long run, domestic nominal shocks affect all domestic variables in X_t in the exception of the real income whereas domestic supply shocks affect all variables in X_t. The results of the estimated common trends model as characterized so far are presented in the following subsection.

3.4 The Estimated Common Trends Model

Based on the information derived above concerning the number of stochastic trends given the lag length as supported by the data set, as well as the characterization of the A_0 matrix we proceed to estimate the parameters and discuss the results of the common trends model in this section. Specifically, we compute the impulse response functions and variance decompositions and utilize these to show how the common trends in the data set affect interest rate differentials, the nominal exchange rate and the current account balance in the countries considered. The parameters of the estimated common trends model are presented in table 3.2. below. The $A_0\pi$ $(=A)$ matrix was chosen so that the elements of the first column is interpreted as domestic real trend ,τ_R, and the second column as domestic nominal shocks ,τ_N. Domestic nominal shocks have zero long run effect on output.

From the results presented above we see that (in the case of Sweden) positive domestic real trends increase the interest rate differential (i.e. i_t^D - i_t^{US}), appreciate the domestic currency vis-a-vis the dollar and worsens the current account balance in the long run. This result is consistent with Kouri(1976) where positive real domestic disturbances appreciate the domestic currency and increase foreign asset accumulations. This result is replicated for both Germany and Japan except for the wrong sign of the coefficient of the interest rate differential. In the case of the UK the wrong sign showed up in the coefficient in front of the interest rate differential and the exchange rate. Domestic nominal trends have no real effects in the long run as predicted by theory. Their effects on the interest rate differential have wrong signs in the cases of all four countries considered. They do however depreciate the domestic currency and improve the current account balance supporting the Marshall-Lerner condition in the cases of all countries in the exception of Japan where the effect on the current account balance has the wrong sign. These are the results of the long run effects as captured by the estimated common trends model.

Table 3.2: Estimates of the Common Trends Models. $X_t = [\ ln\ y_t \quad (i_t^D - i_t^{US}) \quad e_t \quad ca_t\ ]'$.

Germany

$$\begin{bmatrix} 0.0032 & 0.0000 \\ (0.0012) & (-) \\ -0.4434 & 0.3162 \\ (0.2674) & (0.0950) \\ -0.0385 & 0.0323 \\ (0.0252) & (0.0097) \\ -0.3331 & 0.8247 \\ (0.4612) & (0.2478) \end{bmatrix}$$

Japan

$$\begin{bmatrix} 0.1972 & 0.000 \\ (0.0629) & (-) \\ -0.1678 & 0.7903 \\ (0.3755) & (0.2585) \\ -0.2460 & 0.0924 \\ (0.0963) & (0.0302) \\ -3.3000 & -1.4826 \\ (1.1290) & (0.4850) \end{bmatrix}$$

Sweden

$$\begin{bmatrix} 0.0389 & 0.000 \\ (0.0087) & (-) \\ 0.0983 & 1.0016 \\ (0.5060) & (0.4529) \\ -0.0054 & 0.3250 \\ (0.1639) & (0.1469) \\ -0.2059 & 1.1922 \\ (0.6028) & (0.5390) \end{bmatrix}$$

UK

$$\begin{bmatrix} 0.2617 & 0.0000 \\ (0.1944) & (-) \\ -0.0295 & 0.0258 \\ (0.0294) & (0.0058) \\ 0.1506 & 0.0594 \\ (0.1219) & (0.0133) \\ -1.4222 & 0.3015 \\ (1.0775) & (0.0678) \end{bmatrix}$$

Note: *Standard errors are represented in parentheses. The first and second columns of each matrix above correspond to the estimated parameters measuring the long run effects of domestic real and nominal trends respectively.*

Next we review the effects of these shocks in the short-term, medium-term and long-term using the impulse-response functions presented in figures 3.1 - 3.4 as well as the forecast error decompositions shown in tables 3.4 - 3.6.

(i) ***Effects of domestic technological (supply) and nominal shocks***

From our impulse-response figures 3.1 - 3.4 below we see that a one standard deviation shock to real domestic trends appreciate the Deustche mark, the swedish kronor, and the japanese yen - vis-a-vis the US dollar - over the first four quarters, though the results from the data set for Germany are not statistically significant. The effects are however mixed in the long run though the general impression one gets by observing the impulse-responses is that of an appreciation of the domestic currencies. The current account balance worsens over the same period for all cases considered except for the case of Germany where technological shocks lead to improvements in the current account balance over the first fifteen quarters and worsens it afterwards. Again, these results are not statistically significant. Neither is the evidence for Sweden. The positive effect of real domestic shocks on the interest rate differential is felt in the first quarter in both Japan and Sweden. For Germany and the UK the effect of these real shocks on the interest rate differential is negative - we guess here that the model has not being able to adequately capture the transmission mechanism of the shocks on the interest rate differential.

Even though nominal trends seem to depreciate the domestic currencies and improve the current account balances in accordance with the Marshall-Lerner condition they tend to have no significant effects on either the exchange rates or on the current account balances - a rather surprising result.

In the exception of Sweden, for all countries considered, real trends explain 58 - 88 percent of variations in nominal exchange rates whereas nominal trends explain 12 - 40 percent. The corresponding figures for the current account balances are 83-96 percent and 4 - 17 percent respectively in the cases of Japan and the UK. For Sweden and Germany the corresponding figures are 3 - 14 per cent and 85 - 97 per cent for real and nominal trends respectively.

(ii) Effects of permanent and transitory shocks

Within the common trends framework the most interesting results when it comes to innovation accounting are those related to permanent shocks in relation to transitory shocks. Figure 3.4 and Table 3.6 summarize the role of the two transitory and permanent shocks for the four countries in explaining variations in the interest rate differential, nominal exchange rate and the current account balance. We infer from figure 3.3 that, in the cases of Japan and Sweden, transitory real shocks significantly appreciate domestic currencies against the dollar (the effect being felt after five quarters and around seven-eight quarters) and worsen current account balances significantly over the first-seven quarters in Japan and in the third quarter in Sweden.

For all countries considered transitory nominal trends tend to significantly worsen the current account balance over the first-four quarters on the average. These trends explain, according to Table 3.6, roughly 39% - 88% of the total variation in the current account balance over the same period. The exchange rate depreciating effects of these trends are felt over the first-eight quarters in Sweden, and in the fifth/sixth quarter in Japan. Their effect on the interest rate differential is felt during the first-seven quarters in Sweden and in the UK where these shocks lead to reductions in the interest rate differential.

Permanent shocks explain all variations in the exchange rates and the current account balances of the countries considered in the long run. However, over the four quarter period transitory shocks[10] explain roughly 13 - 60 percent and 39 - 88 percent of all variations in the exchange rates and current account balances respectively.

[10] Note that the sum of the two permanent forecast error variance decompositions is 1 minus the sum of the two transitory forecast error variance decompositions.

(iii) ***Effects of fiscal and monetary policy***

From the estimation of the vector autoregressive model using the Johansen routine we derive the estimates of the short run effects of the exogenous variables - the budgetary deficit and the logarithm of money. We can however, only analyze the short-run effects of these variables the effects of which are presented in Table 3.3 below:

The results reveal that only few of the short run coefficients are significant. For Sweden the results indicate that budgetary deficits increase the interest rate differential, appreciate the domestic currency significantly and worsen the current account balance. This is a result that is consistent with Branson(1983 and 1984) and the Marshall-Lerner condition. The significant appreciation of the domestic currency is also evident from the results of Germany even though the (insignificant) effects on the interest rate differential and the current account balance have wrong signs. For Japan and the UK the estimates of the coefficients of the effects of the fiscal policy index (which are all insignificant at the 5% level) seem inconsistent with the theory except for its effect on the current account balance in Japan. For Sweden and the UK monetary policy reduces the interest rate differential, depreciates the domestic currency and improves the current account balance. This result seems consistent with Branson(1983 and 1984), Dornbusch(1976) and Kouri(1976) as well as the Marshall-Lerner condition. However, for Germany and Japan the results are rather unsatisfactory.

Table 3.3: Short-Run Effects of Budget Deficits and Money on $X_t = [ln\, y_t \;\; (i_t^D - i_t^{US}) \;\; e_t \;\; ca_t]'$.

Germany

$$\begin{bmatrix} 0.000 & 0.104 \\ (0.1) & (2.2) \\ -0.030 & 5.822 \\ (0.9) & (1.3) \\ -0.010 & -0.380 \\ (2.7) & (0.7) \\ 0.070 & -41.508 \\ (0.7) & (2.8) \end{bmatrix}$$

Japan

$$\begin{bmatrix} 0.005 & -1.279 \\ (1.1) & (2.9) \\ -0.001 & 0.835 \\ (0.0) & (0.4) \\ 0.026 & -1.288 \\ (1.8) & (1.0) \\ -0.396 & -64.996 \\ (1.8) & (3.3) \end{bmatrix}$$

Sweden

$$\begin{bmatrix} -0.001 & 0.001 \\ (1.9) & (1.9) \\ 0.014 & -0.005 \\ (1.4) & (1.0) \\ -0.008 & 0.004 \\ (3.7) & (3.5) \\ -0.003 & 0.006 \\ (0.2) & (0.7) \end{bmatrix}$$

UK

$$\begin{bmatrix} -0.006 & 0.708 \\ (0.5) & (2.9) \\ -0.027 & -0.138 \\ (0.7) & (0.2) \\ 0.012 & 0.057 \\ (1.8) & (0.4) \\ 0.067 & 6.276 \\ (0.7) & (3.1) \end{bmatrix}$$

Note: *T-values in parenthesis. The first and second columns in each matrix above refer to the effects of budget deficit and of money supply shocks respectively.*

3.5 Summary and Conclusion

This chapter adopts a multivariate data analysis approach - Common Trends Approach - to the analysis of the effects of domestic real and nominal shocks on the nominal exchange rate and the current account balance. We followed the approach of Warne(1990) utilizing the estimation algorithms written by the same author.

We investigated the effects of real and nominal trends on the interest rate differentials (between the US and the four countries, Germany, Japan, Sweden and the UK), exchange rates (defined here as units of domestic currency per US dollar), and the current account balance using time series data on output, budgetary deficits, money supply, interest rate differentials exchange rates and the current account balance for the four countries mentioned above. We found evidence that domestic technological trends (or supply shocks) do have more significant effects on the exchange rate and current account balance over the short and the long run than nominal trends do. Our results also show that transitory shocks affect the exchange rate and the current account balance only *transitorily* (that is to say their effects are felt mainly in the short and medium term). Our short run results of the effects of budgetary deficits and money supply of the exchange rate and the current account balance are consistent with international monetary theory. However the nature of the common trends approach as adopted in this chapter did not permit us to investigate the medium term as well as the long term effects of the budgetary deficit and monetary policy on the exchange rate and the current account balance. This issue will be taken up in another project.

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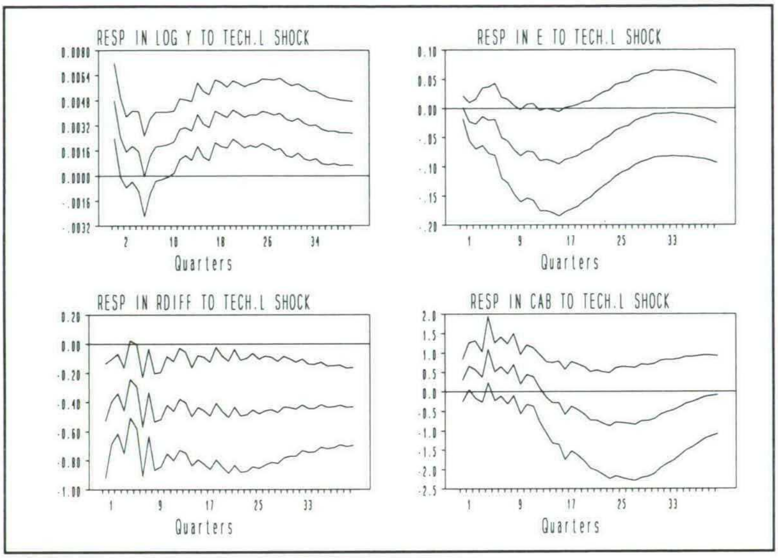
Figure 3.1: Responses in X_t following a one-standard deviation shock in the domestic real trend (τ_R) with 95% confidence bounds

Germany



Japan

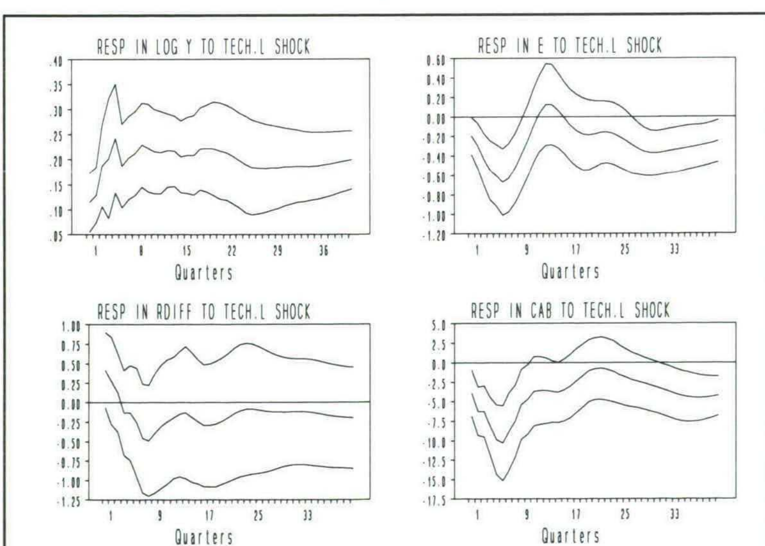
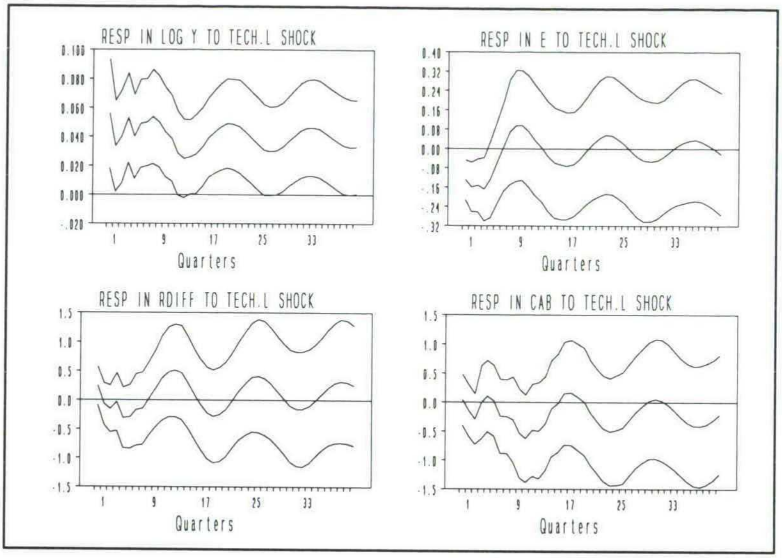


figure 3.1 contd..

Sweden



UK

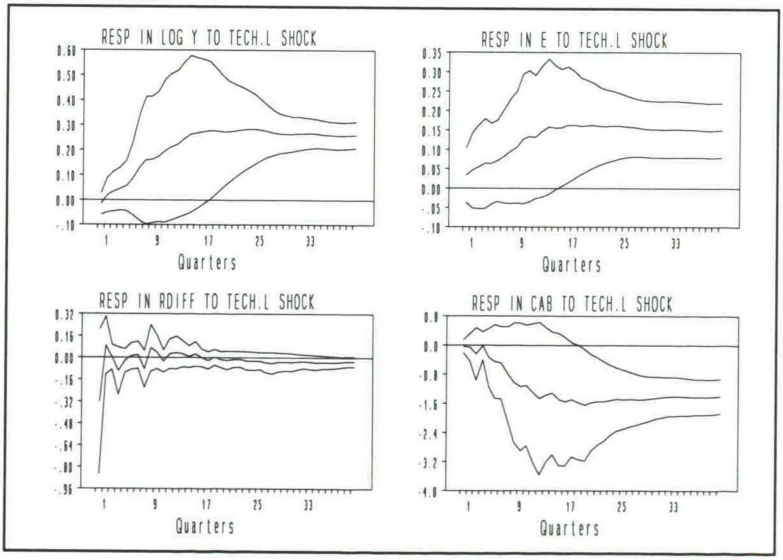
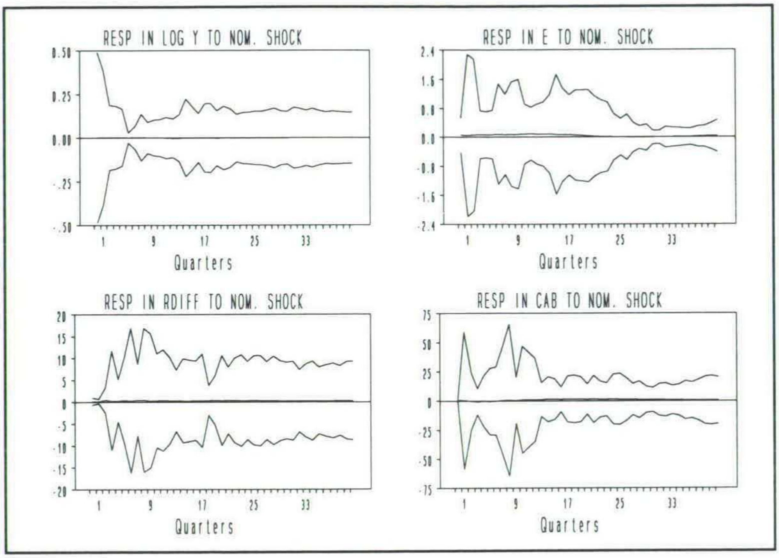


Figure 3.2: Responses in X_t following a one-standard deviation shock in the domestic nominal trend (τ_N) with 95% confidence bounds

Germany



Japan

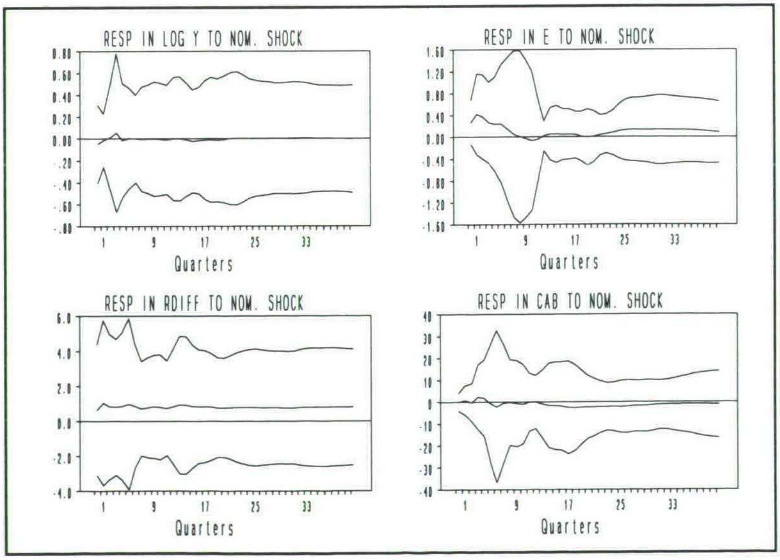
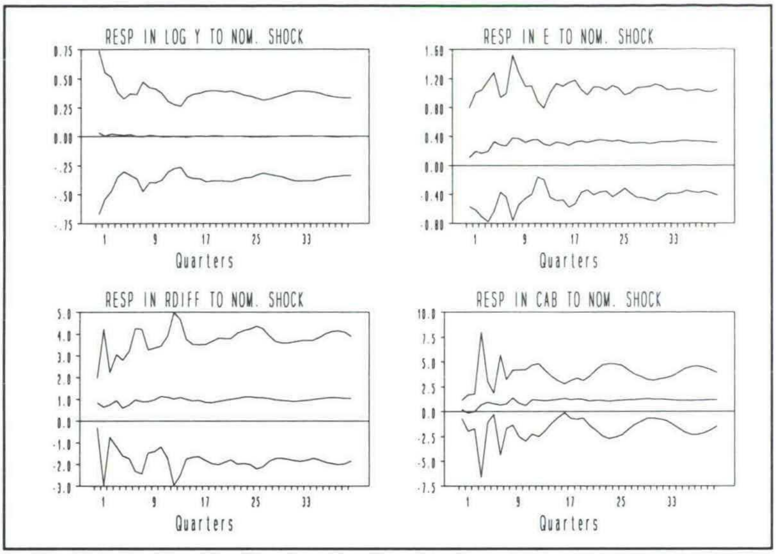


figure 3.2 contd.

Sweden



UK

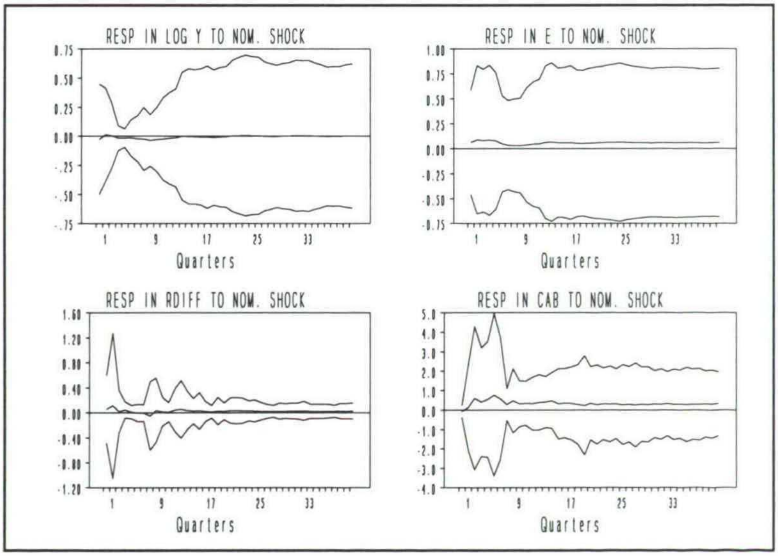
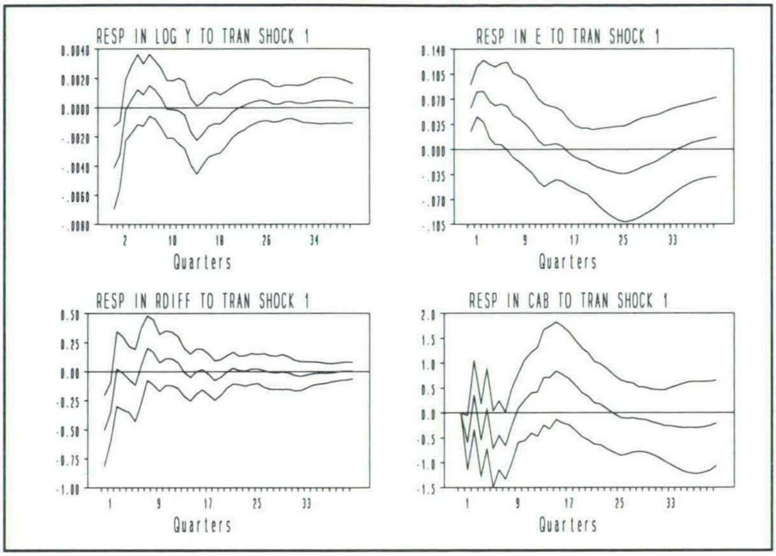


Figure 3.3: Responses in X_t following a one-standard deviation transitory shock in the domestic real trend (τ_R) with 95% confidence bounds

Germany



Japan

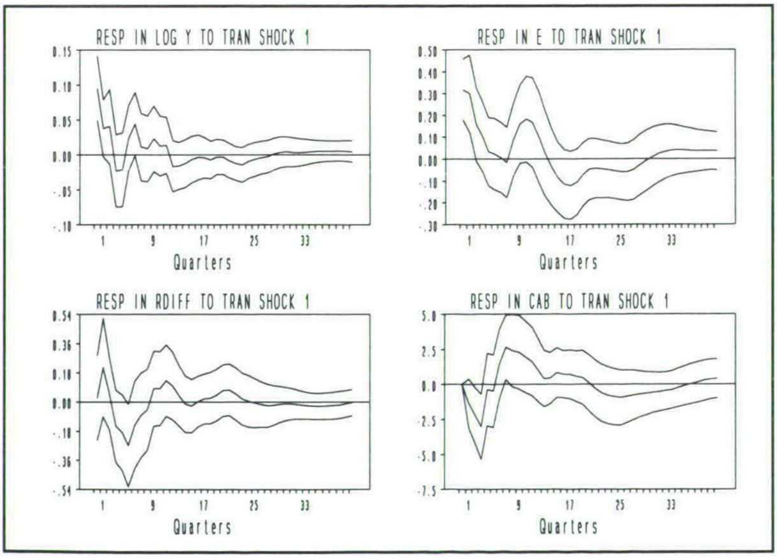
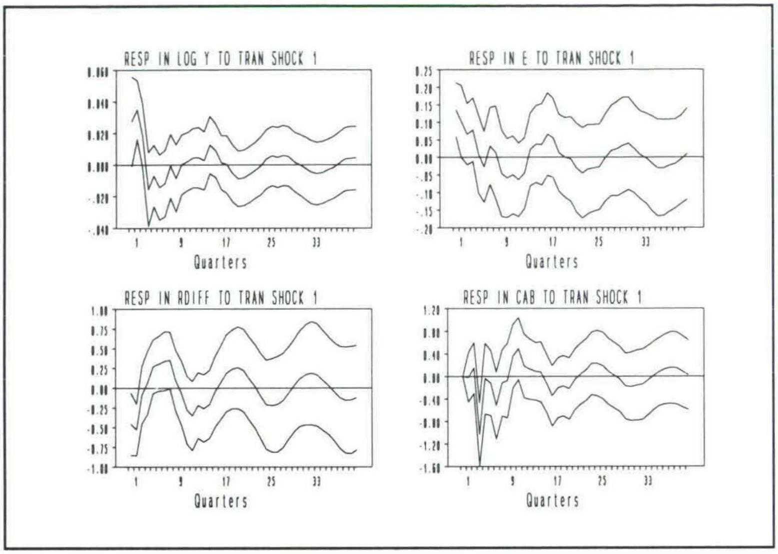


figure 3.3 contd.

Sweden



UK

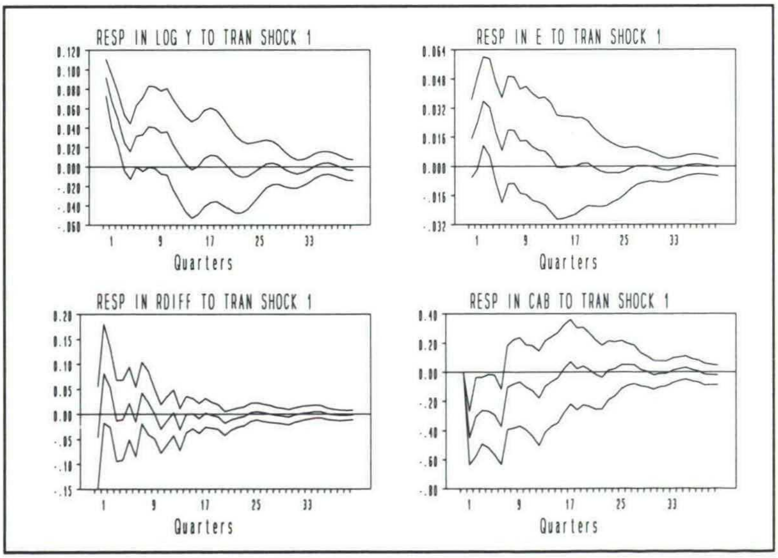
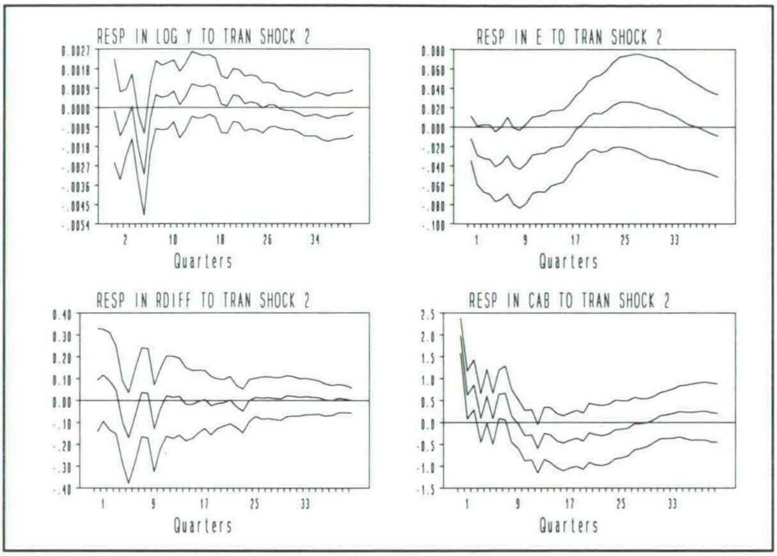


Figure 3.4: Responses in X_t following a one-standard deviation transitory shock in the domestic nominal trend (τ_N) with 95% confidence bounds

Germany



Japan

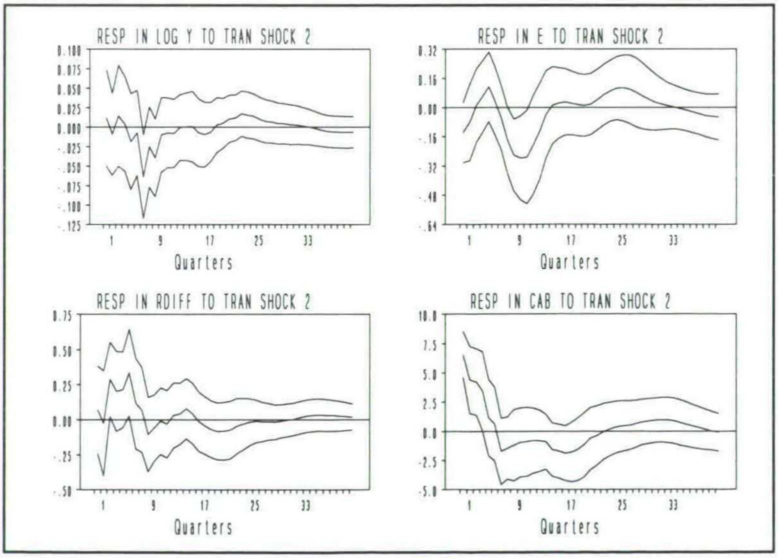
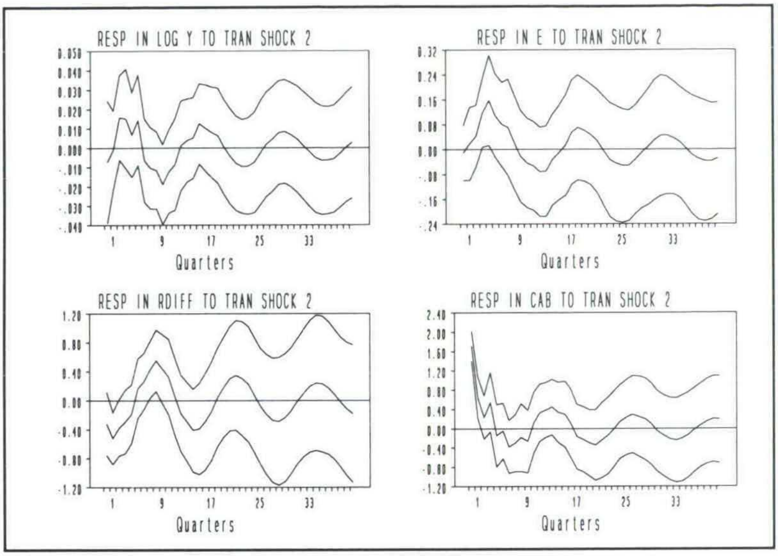


figure 3.4 contd.

Sweden



UK

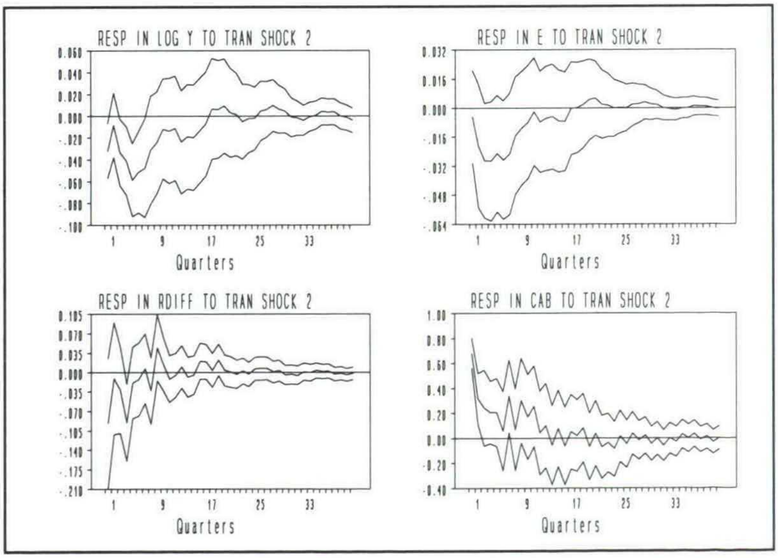


Table 3.4: Ratio of s steps ahead Forecast error Variance of X_t accounted for by the domestic real trend (τ_R)

	Germany				Japan			
s	*ln y*	r^p	*e*	*ca*	*ln y*	r^p	*e*	*ca*
1	0.530 (-)	0.497 (-)	0.000 (-)	0.019 (-)	0.537 (0.001)	0.290 (-)	0.168 (0.023)	0.263 (-)
2	0.427 (-)	0.499 (-)	0.032 (0.003)	0.093 (-)	0.695 (0.002)	0.138 (-)	0.216 (0.196)	0.458 (-)
3	0.412 (-)	0.461 (-)	0.046 (0.013)	0.120 (-)	0.813 (0.004)	0.103 (-)	0.337 (0.632)	0.519 (-)
10	0.355 (-)	0.537 (-)	0.166 (0.446)	0.252 (-)	0.930 (0.024)	0.127 (-)	0.683 (9.248)	0.785 (-)
20	0.569 (-)	0.581 (-)	0.359 (3.339)	0.146 (-)	0.964 (0.043)	0.102 (-)	0.645 (17.56)	0.754 (-)
30	0.713 (-)	0.608 (-)	0.380 (5.627)	0.199 (-)	0.973 (0.048)	0.080 (-)	0.669 (31.49)	0.723 (-)
40	0.756 (-)	0.623 (-)	0.377 (7.326)	0.204 (-)	0.978 (0.057)	0.074 (-)	0.708 (55.61)	0.753 (-)
∞	1.000 (-)	0.663 (0.262)	0.586 (0.303)	0.140 (0.307)	1.000 (-)	0.043 (0.182)	0.876 (0.106)	0.832 (0.14)

	Sweden				UK			
s	*ln y*	r^p	*e*	*ca*	*ln y*	r^p	*e*	*ca*
1	0.628 (0.014)	0.057 (-)	0.350 (-)	0.000 (-)	0.025 (-)	0.885 (-)	0.218 (-)	0.003 (-)
2	0.578 (0.001)	0.032 (-)	0.343 (-)	0.007 (-)	0.038 (-)	0.774 (-)	0.201 (0.003)	0.007 (-)
3	0.591 (0.001)	0.032 (-)	0.365 (0.176)	0.031 (-)	0.085 (-)	0.752 (-)	0.231 (0.005)	0.048 (-)
10	0.773 (0.007)	0.037 (-)	0.136 (2.017)	0.052 (-)	0.729 (0.027)	0.699 (-)	0.570 (0.040)	0.486 (-)
20	0.822 (0.009)	0.069 (-)	0.078 (7.064)	0.061 (-)	0.941 (0.083)	0.661 (-)	0.789 (0.133)	0.823 (-)
30	0.855 (0.010)	0.065 (-)	0.056 (12.81)	0.066 (-)	0.969 (0.099)	0.644 (-)	0.825 (0.240)	0.885 (-)
40	0.883 (0.012)	0.058 (-)	0.044 (18.38)	0.063 (-)	0.978 (0.109)	0.641 (-)	0.836 (0.357)	0.905 (-)
∞	1.000 (-)	0.009 (0.098)	0.000 (0.017)	0.029 (0.167)	1.000 (-)	0.567 (0.508)	0.866 (0.194)	0.957 (0.07)

Note: *The estimated standard errors are based on Theorem 2.4. of Warne (1990). Very insignificantly small or zero standard errors are indicated with (-) in the table.*

Table 3.5: Ratio of s steps ahead Forecast error Variance of X_t accounted for by the domestic nominal trend (τ_N)

	Germany				Japan			
s	*ln y*	r^P	*e*	*ca*	*ln y*	r^P	*e*	*ca*
1	0.079	0.026	0.468	0.066	0.089	0.700	0.315	(-)
	(-)	(-)	(0.008)	(-)	(0.002)	(-)	(0.047)	(-)
2	0.144	0.041	0.316	0.067	0.056	0.833	0.419	0.005
	(-)	(-)	(0.067)	(-)	(0.004)	(-)	(0.414)	(-)
3	0.196	0.197	0.296	0.091	0.031	0.843	0.407	0.004
	(-)	(-)	(0.191)	(-)	(0.005)	(-)	(0.995)	(-)
10	0.283	0.301	0.419	0.161	0.014	0.811	0.174	0.022
	(-)	(-)	(1.15)	(-)	(0.024)	(-)	(9.799)	(-)
20	0.177	0.326	0.424	0.470	0.008	0.861	0.161	0.054
	(-)	(-)	(5.446)	(-)	(042)	(-)	(18.91)	(-)
30	0.119	0.329	0.383	0.555	0.006	0.891	0.155	0.095
	(-)	(-)	(9.188)	(-)	(0.043)	(-)	(34.53)	(-)
40	0.102	0.327	0.384	0.558	0.005	0.907	0.149	0.090
	(-)	(-)	(12.01)	(-)	(0.047)	(-)	(61.47)	(-)
∞	0.000	0.337	0.414	0.859	0.000	0.957	0.124	0.168
	(-)	(0.262)	(0.303)	(0.307)	(-)	(0.181)	(0.106)	(0.14)

	Sweden				UK			
s	*ln y*	r^P	*e*	*ca*	*ln y*	r^P	*e*	*ca*
1	0.209	0.651	0.266	0.014	0.062	0.025	0.732	0.013
	(0.018)	(-)	(-)	(-)	(0.001)	(-)	(-)	(-)
2	0.146	0.540	0.418	0.019	0.052	0.107	0.721	0.028
	(0.002)	(-)	(0.277)	(-)	(0.001)	(-)	0.004)	(-)
3	0.145	0.598	0.437	0.018	0.041	0.106	0.646	0.285
	(0.003)	(-)	(0.575)	(-)	(0.001)	(-)	(0.008)	(-)
10	0.074	0.720	0.753	0.485	0.031	0.113	0.344	0.294
	(0.008)	(-)	(3.363)	(-)	(0.013)	(-)	(0.046)	(-)
20	0.052	0.766	0.846	0.692	0.008	0.150	0.185	0.116
	(0.011)	(-)	(12.99)	(-)	(0.059)	(-)	(0.138)	(-)
30	0.039	0.805	0.887	0.758	0.004	0.178	0.160	0.081
	(0.011)	(-)	(24.75)	(-)	(0.075)	(-)	(0.245)	(-)
40	0.031	0.834	0.911	0.799	0.003	0.192	0.153	0.070
	0.014)	(-)	(37.15)	(-)	(0.083)	(-)	(0.362)	(-)
∞	(-)	0.990	0.999	0.971	(-)	0.433	0.134	0.043
	(-)	(0.098)	(0.017)	(0.167)	(-)	(0.507)	(0.194)	(0.06)

Note: *The estimated standard errors are based on Theorem 2.4. of Warne (1990). Very insignificantly small or zero standard errors are indicated with (-) in the table.*

Table 3.6: Ratio of s steps ahead Forecast error Variance of X_t accounted for by the two permanent shocks

	Germany				Japan			
s	*ln y*	r^P	*e*	*ca*	*ln y*	r^P	*e*	*ca*
1	0.069 (-)	0.523 (-)	0.468 (0.008)	0.088 (-)	0.626 (0.001)	0.990 (-)	0.483 (0.036)	0.263 (-)
2	0.571 (-)	0.540 (-)	0.348 (0.064)	0.016 (-)	0.751 (0.002)	0.971 (-)	0.635 (0.292)	0.463 (-)
3	0.608 (-)	0.658 (-)	0.342 (0.178)	0.211 (-)	0.844 (0.003)	0.946 (-)	0.744 (0.506)	0.523 (-)
10	0.618 (-)	0.838 (-)	0.585 (1.079)	0.413 (-)	0.944 (0.009)	0.938 (-)	0.857 (2.044)	0.807 (-)
20	0.746 (-)	0.907 (-)	0.783 (2.077)	0.616 (-)	0.972 (0.013)	0.963 (-)	0.806 (6.345)	0.808 (-)
30	0.832 (-)	0.937 (-)	0.763 (3.603)	0.754 (-)	0.979 (0.021)	0.971 (-)	0.824 (11.19)	0.818 (-)
40	0.858 (-)	0.950 (-)	0.761 (4.742)	0.762 (-)	0.983 (0.026)	0.981 (-)	0.857 (15.90)	0.843 (-)
∞	1.000 (-)	1.000 (-)	1.000 (-)	1.000 (-)	1.000 (-)	1.000 (-)	1.000 (-)	1.000 (-)

	Sweden				UK			
s	*ln y*	r^P	*e*	*ca*	*ln y*	r^P	*e*	*ca*
1	0.837 (0.004)	0.708 (-)	0.616 (-)	0.014 (-)	0.087 (0.001)	0.910 (-)	0.950 (-)	0.016 (-)
2	0.724 (-)	0.572 (-)	0.761 (0.106)	0.026 (-)	0.090 (-)	0.881 (-)	0.922 (0.001)	0.035 (-)
3	0.736 (0.001)	0.630 (-)	0.802 (0.186)	0.049 (-)	0.126 (-)	0.858 (-)	0.877 (0.003)	0.333 (-)
10	0.847 (0.002)	0.757 (-)	0.889 (1.474)	0.537 (-)	0.760 (0.018)	0.812 (-)	0.914 (0.007)	0.780 (-)
20	0.874 (0.002)	0.835 (-)	0.924 (6.278)	0.753 (-)	0.949 (0.051)	0.811 (-)	0.974 (0.005)	0.939 (-)
30	0.894 (0.002)	0.870 (-)	0.943 (12.51)	0.824 (-)	0.973 (0.051)	0.822 (-)	0.985 (0.005)	0.966 (-)
40	0.914 (0.002)	0.892 (-)	0.955 (19.63)	0.862 (-)	0.981 (0.051)	0.833 (-)	0.989 (0.005)	0.975 (-)
∞	1.000 (-)	1.000 (-)	1.000 (-)	1.000 (-)	1.000 (-)	1.000 (-)	1.000 (-)	1.000 (-)

Note: The estimated standard errors are based on Theorem 2.4. of Warne (1990). Very insignificantly small or zero standard errors are indicated with (-) in the table.

Chapter 4

Common Stochastic Trends in the Current Account

4.1 Introduction

Using the intertemporal approach to the current account Glick and Rogoff(1995) - referred to as G&R in the rest of this paper - develop and empirically test specifications for investment and the current account especially with respect to the effects of global and domestic (country-specific) productivity shocks on these using time series data for the G-7 countries. The framework is basically in the tradition of Obstfeld(1986) and Razin(1993) who analyzed the theoretical effects of government spending and productivity shocks. The main departure in G&R (just as in Razin (1993)) is an empirical re-formulation of the problem to distinguish between global and country-specific shocks. Using this framework G&R investigate the relative effects of permanent global and country-specific productivity shocks on investment and the current account.

Earlier attempts at investigating the effects of productivity shocks and/or stochastic trends on economic fluctuations in an open economy use simulations and/or Blanchard-Quah type vector-autoregression methodology: Mendoza(1991) falls into the former category whereas Ahmed, Ickes, Wang and Yoo(1993) falls into the latter group. There is also a group of researchers who use Solow residuals as proxies for productivity shocks like G&R and Backus, Kehoe and Kydland(1992) for example. The use of Solow residuals as proxies for productivity shocks is however not without shortcomings. Firstly, as pointed out in Hall(1988) movements in the Solow residuals may not represent exogenous technology shocks - the identifying assumption here is for the Solow residuals to be orthogonal to "variables known to be neither causes of productivity shifts nor to be caused by productivity shifts". G&R do not test for this identifying assumption. Secondly, the use of constructed Solow residuals as productivity shocks has been shown to overestimate the variance of these shocks due to increased variation in the capital stock as the result of varying capacity utilization and other measurement problems [see for instance Eichenbaum(1990)]. Baxter and Crucini(1993) is very eloquent on the possibility of these measurement errors carrying over into obtained empirical results when they criticise the mode of measurement of the Solow residuals by Backus *et al*(1992).

This paper adds to the results in G&R by adopting multivariate cointegration and common trends techniques that circumvent the measurement problems associated with calculated Solow residuals. The idea here is to decompose the variables of interest directly into permanent and transitory components rather than search for the response of these variables to calculated Solow residuals. Moreso, given that the number of cointegrating vectors is different from zero we can be sure there are certain common factor(s) that explain fluctuations in the variables of interest. We use the common trends technique to identify and interpret these factors as productivity shocks. Our focus here is on the responses of investment and the current account balance to country-specific and global productivity shocks. The specific questions that we seek empirical answers to are the following: i. what are the long run effects of each of these shocks (analyzed separately) on investment and the current account? ii. how does each of our variables of interest (investment and the current account) respond to one standard deviation innovations in the respective shocks? iii. what is the relative effectiveness of the innovations in the respective shocks in explaining fluctuations in investment and the current account? and finally iv. how does our approach compare with that of G&R? In an attempt to address these questions we utilize common trends approach in identifying and estimating the effects of these shocks following the estimation structure as presented in Johansen(1988), King *et al.*(1991), Stock and Watson(1988) and Warne(1990). Our empirical investigations are based on data-sets for Germany and the United States (US) over the period 1974:1-1992:4.

Our empirical results indicate that the common trends approach yields results/findings comparable to those of G&R and goes beyond instantaneous least squares estimates by providing us with the possibility of dynamic analyses - which in this context could be implemented using estimated impulse response functions of the effects of innovations in the identified productivity shocks and forecast error decompositions - that are very crucial in empirical investigations of capital mobility and current account fluctuations. It turns out that the estimated productivity shocks are highly persistent and explain almost all variations in our variables of interest at the long run horizon. In fact the estimated impulse response functions attributable to innovations in the transitory components of these shocks - as portrayed by figures 4.4 and 4.5 - show that they do not seem relevant in explaining fluctuations in our data-set. Also, long run fluctuations in real output and consumption are better explained by global shocks whereas fluctuations in investment and the current account are better accounted for by idiosyncratic productivity shocks. This evidence is indicative of the fact that the two countries considered are *very* open economies in which capital is highly mobile and capital flights are carried out in attempts to smoothen consumption - the usual risk-sharing arguments - and hence making consumption highly correlated with global shocks. Given this explanation agents consume a share of the world consumption/output irrespective of domestic idiosyncratic disturbances.

In pursuit of answers to the specific questions raised above we organise the paper as follows. Section 4.2 presents the theoretical framework used in deriving the reduced-form equations for investment and the current account and distinguishes carefully between the effects of shocks, that could be country-specific or global. We adopt the specification of G&R and Razin(1993). Theoretical implications of the model are derived and discussed in this section. Section 4.3 inspects the data-set for seasonal integration and tests for the number of cointegrating vectors (or equivalently, the number of common trends), if any. The section after that presents and analyses the estimates of the common trends model the main features of which are summarized in chapter 1 of this dissertation. The estimates of the impulse response functions are also presented here. Finally, section 4.5 summarises and concludes the chapter.

4.2 Theoretical Framework

In the tradition of the dynamic-optimizing approach[1] to the current account balance fluctuations in the current account are attributable, among other factors, to productivity shocks (which could be transitory or permanent) characterised as global or country-specific and are transmitted through savings-investment decisions of economic agents.

4.2.1 The Investment Decision

Consider an economy producing a single aggregate tradable good, Y_t, using the Cobb-Douglas production function in (4.1) below. We assume here that labour is supplied inelastically such that

$$Y_t = Z_t K_t^{\alpha} \tag{4.1}$$

where Z_t denotes the time-t productivity, K_t is the capital stock and α is the distributive share of capital. Following G&R we assume that Z_t represents the time - t stochastic shock to technology and that it follows a first-order autoregressive process given by

$$Z_t = \rho Z_{t-1} + \epsilon_t, \qquad where \qquad \rho \in [0,1] \tag{4.2}$$

where ρ is a persistence parameter[2] and ε_t is a white-noise disturbance term. The representative firm maximises the expected value of the discounted sum of profits

[1] See for instance Razin(1993), Obstfeld(1986) and/or G&R for an exposition of this approach.

[2] We solve the model for $\rho \in [\ 0,1\ ]$ though our interest is best served with $\rho = 1$ (a random-walk specification for the productivity term.

$E\sum_{t=0}^{\infty} R^{-t}\pi_t$ where $\pi_t = Y_t - \omega_t$, and ω_t denotes the cost-of-adjustment of investment technology

specified as

$$\omega_t = I_t \left[1 + \frac{g}{2}\frac{I_t}{K_t} \right] \qquad (4.3)$$

where $K_{t+1} = K_t + I_t$ and g is the cost-of-adjustment coefficient. Given this cost-of-adjustment technology there is an incentive for firms to adjust their capital stocks gradually since the cost of changing the capital stock by one unit increases with the speed of adjustment. Using Euler equations the optimal investment rule is characterised as below:

$$E_t R^{-1} \left[\alpha Z_{t+1} K_{t+1}^{\alpha-1} + \frac{g}{2}\left(\frac{I_{t+1}}{K_{t+1}}\right)^2 + q_{t+1} \right] = q_t \qquad (4.4)$$

where E_t is the expectations operator based on the current period's information, q_t is the firm's market-value per unit of capital such that $q_t = 1 - \frac{g}{2}\frac{I_t}{K_t}$ and R is one plus the world rate of interest. Notice that linearising (4.4) around the steady-state (where there is equality between the world rate of interest and the marginal productivity of capital: i.e $R - 1 = \alpha \overline{Z}\overline{K}^{\alpha}$ where $\overline{Z}$ and $\overline{K}$ are the steady state levels of productivity and capital,stock respectively) yields the expression below (where the lag operator, L, is defined as $L^k(\chi_{t+1}) = E_t(\chi_{t+1-k})$ for both positive and negative values of k)

$$a_1 \left[L^0 + \phi_1 L + \phi_2 L^2 \right] k_{t+1} = - bZ_t \qquad (4.5)$$

where $k_t = K_t - \overline{K}$ denotes the deviation of capital stock from its steady state value, $\phi_2 = (1/a_1)$ and $\phi_1 = (a_0/a_1)$. The parameters a_0 and a_1 are functions of R, g and α. The polynomial on the left-hand side of the above expression can be factorised (see Sargent(1979)) as

$$\left[L^0 + \phi_1 L + \phi_2 L^2\right] = (L^0 - \varphi_1 L)(L^0 - \varphi_2 L)$$
$$= -\varphi_2 L(L^0 - \varphi_2^{-1} L^{-1})(L^0 - \varphi_1 L)$$

We assume $\phi_1 + \phi_2 < 1$, $\phi_2 - \phi_1 < 1$, and $\phi_2 < 1$ so that $(L^0 - \varphi_2^{-1}L^{-1})$ is invertible. [see for instance Cryer(1989) for the derivation of these conditions]. Multiplying both sides of (4.5) by the inverse of $-a_1\varphi_2 L(L^0 - \varphi_2^{-1}L^{-1})$ yields

$$k_t = \lambda_1 k_{t-1} + \lambda_1 b \sum_{i=0}^{\infty}(\frac{1}{\lambda_2})^i E Z_{t+1+i} \qquad (4.6)$$

where $\lambda_1 < 1$ and $\lambda_2 > 1$ are the roots of the quadratic equation $1 + a_0\lambda + a_1\lambda^2 = 0$. A simple manipulation of equation (4.6) - neglecting the i.i.d term in the AR(1) productivity shock process - yields desired investment as

$$I_t = \lambda_1 I_{t-1} + \lambda_1 b \sum_{i=1}^{\infty}(\frac{1}{\lambda_2})^{i-1}\left[E_t\rho^i Z_{t+i} - E_{t-1}\rho^i Z_{t+i-1}\right]$$

$$= \lambda_1 I_{t-1} + \lambda_1 b\rho \sum_{i=1}^{\infty}(\frac{\rho}{\lambda_2})^{i-1}\Delta Z_t$$

$$= \lambda_1 I_{t-1} + \frac{\lambda_1 b\rho}{(1 - \frac{\rho}{\lambda_2})}\Delta Z_t \qquad (4.7)$$

$$\Rightarrow \Delta I_t = (\lambda_1 - 1)I_{t-1} + \frac{\lambda_1 b\rho}{(1 - \frac{\rho}{\lambda_2})}\Delta Z_t$$

The first term of the first line of the above equation captures the effects on current investment of lagged productivity shocks and the second term the revisions in expectations of future productivity shocks. Thus transitory productivity shocks have no impact on current investment.

4.2.2 Consumption

The representative agent chooses a consumption path that maximizes the lifetime utility

$$E_t\left[\sum_{i=1}^{\infty}\beta^i U(C_{t+i})\right], \qquad \textit{where} \quad U(c) = C - \frac{1}{2}C^2 \tag{4.8}$$

subject to the intertemporal budget constraint

$$F_t = Y_t - C_t + RF_{t-1} \tag{4.9}$$

where β and F denote the subjective discount factor and the stock of foreign assets respectively. Assuming for simplicity that $\beta R \approx 1$, the solution to this optimisation problem can be expressed compactly as

$$C_t = \delta W_t, \qquad \textit{for} \quad \delta = \frac{R-1}{R} \tag{4.10}$$

where W_t denotes the permanent net (investment) income at time t consisting of the expected discounted flow of current and future income and initial foreign assets:

$$W_t = E_t\left[\sum_{i=0}^{\infty}R^{-i}Y_{t+i}\right] + RF_{t-1} \tag{4.11}$$

It is clear that the induced optimal investment path and hence the realised current and future productivity shocks are the main driving forces behind consumption spending according to this model. To ascertain this we linearise the production function around the steady state yielding which is then substituted together with equations (4.2) and (4.7) into the wealth term in equations (4.11) and (4.10) yielding the closed-form solution for optimal consumption as

$$Y_t = d_0 + d_K K_t + d_Z Z_t \tag{4.12}$$

dependent on past and current productivity shocks and the level of initial foreign asset holdings. Specifically, after the necessary substitutions have been made the first difference of equation (4.12) can be expressed as

$$\begin{aligned} \Delta Y_t &= d_K \lambda_1 I_{t-1} + \frac{d_K \lambda_1 b \rho}{(1 - \frac{\rho}{\lambda_2})} \Delta Z_t + d_K \Delta Z_t \\ &= d_K \lambda_1 I_{t-1} + \psi \Delta Z_t \end{aligned} \tag{4.13}$$

Notice here that $\psi > 0$ (for $\lambda_2 > \rho$) indicating that positive technology shocks have positive effects on output. Some further algebraic manipulations then give us the closed-form solution of the first difference of consumption

$$\begin{aligned} \Delta C_t &= \delta \Delta W_t \\ &= \delta E_t \left[\sum_{j=0}^{\infty} R^{-j} \Delta Y_{t+j} \right] + \delta (R - 1) \Delta F_{t-1} \\ &= \frac{\delta d_K \lambda_1}{(1 - \frac{1}{R})} I_{t-1} + \frac{\delta \psi}{(1 - \frac{1}{R})} \Delta Z_t + \delta R \, \Delta F_{t-1} \end{aligned} \tag{4.14}$$

Armed with equations (4.7), (4.13) and (4.14) we derive the equilibrium expression for the change in the current account using the national income accounting identity

$$\Delta ca_t = \Delta Y_t - \Delta I_t - \Delta C_t + (R - 1)\Delta F_{t-1} \tag{4.15}$$

Substituting equations (4.7), (4.13) and (4.14) for the terms on the right hand side of the above expression yields the first difference of the current account as a function of changes in foreign asset holdings and productivity shocks as expressed below:

$$\begin{aligned}
\Delta ca_t &= d_K\lambda_1 I_{t-1} + \psi\Delta Z_t - (\lambda_1 - 1)I_{t-1} - \frac{\lambda_1 b\rho}{(1 - \frac{\rho}{\lambda_2})}\Delta Z_t \\
&\quad - \frac{\delta d_K \lambda_1}{(1-\frac{1}{R})} I_{t-1} - \frac{\delta\psi}{(1 - \frac{1}{R})}\Delta Z_t + [(R - 1) - \delta R]\Delta F_{t-1} \\
&= [\ \lambda_1(d_K - \frac{d_K\delta}{(1-\frac{1}{R})} - 1) + 1\ ]I_{t-1} + [(R - 1) - \delta R]\Delta F_{t-1} \\
&\quad + \left[\psi - \frac{\lambda_1 b\rho}{(1 - \frac{\rho}{\lambda_2})} - \frac{\delta\psi}{(1 - \frac{1}{R})}\right]\Delta Z_t
\end{aligned} \tag{4.16}$$

Thus, if the coefficient in front of ΔZ_t is negative, permanent country-specific productivity-raising shocks must worsen the current account balance. The explanation for this is simply the fact that these shocks, as we discussed above, do not cause only investment spending to rise but also do cause consumption spending to rise at least by as much as the rise in output emanating from these same shocks. These are the issues that we investigate in the empirical section of the paper.[3]

[3] Those interested in the technical details of the econometric methods used here can turn to chapter 1 of this dissertation for a brief exposition on common trends and related issues. Otherwise the reader can move on to the next section without losing the thread of the discussion.

4.3 Preliminary Data Analyses

4.3.1 The Data

The data used in the empirical analysis include seasonally adjusted quarterly data on real gross national product (y_t), private consumption (c_t), gross investment (inv_t) and the current account balance (ca_t) obtained from International Financial Statistics (IFS) database as published by the IMF. The choice of the variables in the data-set is guided by the main variables of the intertemporal approach to the current account as presented in the theoretical section above. The other relevant variables - the world interest rate and net foreign asset holdings - are assumed to be already captured by the current account which by definition is the sum of net exports and interest earnings on net foreign asset holdings (i.e. the product of the world interest rate and the net foreign asset holdings). Since the quarterly current account series are reported/expressed in US dollars we convert them into deutsche marks (in the case of Germany) by multiplying by the average market exchange rates for the respective quarters. Following G&R investment is defined as gross fixed capital formation plus changes in (inventory) stocks. Consumption is defined as private consumption expenditures.

4.3.2 Seasonal Integration

It is quite advantageous when dealing with seasonal data to start by examining the set of plots as depicted in Figures 4.1a and 4.1b for Germany and the US respectively. The first row of each figure shows the level of the series (in the first column) and plots the first quarter values of the series, q_1, the second quarter values, q_2, and so on (in the second column). Thus the quarterly series are graphed in yearly terms. According to Bowswijk and Franses(1991) the plots of these q_i series will be parallel to each other if the seasonal movements in the data are constant (and hence can be satisfactorily modelled using dummies) whereas for a varying seasonal movement (which is better modelled by a stochastic model) they are non-parallel. The last columns of the first row as well as the second row show plots of transformations of the original series based on the transformation

$$(1 - L^4) = (1 - L)(1 + L + L^2 + L^3) = (1 - L)(1 + L)(1 + L^2)$$
$$= (1 - L)(1 + L)(1 - iL)(1 + iL)$$

where L is the lag operator and $i^2 = -1$. Hence if the transformation above renders the quarterly series stationary then the quarterly seasonal unit root process has four roots of modulus unity: one at the zero frequency (which can be removed using the transformation $(1 - L)$), one at the two-quarter or half-yearly frequency (which can be removed using the transformation $(1 + L)$) and a pair of complex conjugate roots at the four-quarter or yearly frequency as captured by $(1 - iL)(1 + iL)$. The figures titled "*NO ROOTS*", "*ZERO-FREQUENCY ROOT*", "*SEMI-ANNUAL ROOT*" and "*ANNUAL ROOT*" depict these respective transformations. It does seem reasonable, judging from these plots, to conclude that generally speaking seasonality in the data can be satisfactorily modelled using seasonal dummies.

This hypothesis of constant seasonal movements as opposed to that of a varying seasonal pattern in the data is formally testable using the testing strategy proposed by Hylleberg *et al.* (1990) - the HEGY procedure. The test procedure requires OLS estimation of the equation

$$\Delta_4 X_t = \pi_1 z_{1t-1} + \pi_2 z_{2t-1} + \pi_3 z_{3t-2} + \pi_4 z_{3t-1} + \epsilon_t \qquad (4.17)$$

and the estimated value of the πs used to draw inferences. In the above equation X_t is the original series, $z_{1t} = (1 + L + L^2 + L^3)X_t$, $z_{2t} = -(1 - L + L^2 - L^3)X_t$ and $z_{3t} = -(1 - L^2)X_t$. Lags of the dependent variable, $\Delta_4 x_t$ could be added to capture autocorrelation in the error term. To test the null hypothesis of a unit root at the zero-frequency we simply test $\pi_1 = 0$; to test for a root of -1 (half-yearly frequency) we test $\pi_2 = 0$ and finally to test for roots of

$\pm i$ (annual frequency) we perform the joint test $\pi_3 = \pi_4 = 0$. If none of the null hypotheses above can be accepted then the original series is stationary. Critical values of these null hypotheses against their respective alternatives $\pi_1 < 0$, $\pi_2 > 0$ and $\pi_3 \cup \pi_4 \neq 0$ are taken from Hylleberg *et al.*(1990). The results of this HEGY testing strategy, applied to the data set, are presented in Table 4.1 below. An intercept term, three seasonal dummies and a linear trend are included in all the regressions performed except in the case of the ca_t where an additional test is conducted with no trend included in the regressions since (from a pragmatic point of view it may be more appropriate to regard the ca_t series as non-trending otherwise current account imbalances will be self-sustaining. However - as the results show - the inclusion, or otherwise, of the trend term in this case (as well as in the case of the other variables) does not yield any qualitative differences in the results.

Table 4.1: Results of the HEGY tests

A. Germany

Series	Augm.	$t(\pi_1)$	$t(\pi_2)$	$F(\pi_3 \cap \pi_4)$	BP(30)	ARCH(1)	ARCH(4)
y_t	0	-3.486	-3.314*	5.700*	0.998	0.793	0.707
c_t	0	-0.596	-4.954*	36.033*	0.109	0.847	0.718
inv_t	0	-1.926	-3.561*	58.373*	0.707	0.875	0.995
ca_t	0 [0]	-2.509 [-1.732]	-3.036* [-3.024*]	43.176* [41.080*]	0.476 [0.495]	0.961 [0.969]	0.852 [0.902]

B. US

Series	Augm.	$t(\pi_1)$	$t(\pi_2)$	$F(\pi_3 \cap \pi_4)$	BP(30)	ARCH(1)	ARCH(4)
y_t	0	-2.701	-4.969*	27.177*	0.382	0.966	0.999
c_t	0	-2.217	-4.716*	34.059*	0.879	0.886	0.822
inv_t	0	-2.499	-6.178*	22.641*	0.967	0.757	0.754
ca_t	1 [1]	-0.672 [-1.527]	-0.710 [-0.724]	0.885 [0.989]	0.963 [0.963]	0.642 [0.632]	0.229 [0.207]

Notes: *Augm. (=Augmentation) depicts the number of lags of the dependent variable included in the regression to attain i.i.d. residuals. P.Values appear under each of the columns labelled 'BP(30)', 'ARCH(1)', and 'ARCH(4)'. 'BP(30)' is the Box-Pierce test for residual autocorrelation based on 30 correlations whereas 'ARCH(k)' tests for autoregressive conditional heterogeneity, at lag k, in the residuals. Rejection of the null hypotheses at the 5% and 10% significant levels are indicated with '*' and '**' respectively. The critical values are taken from Hylleberg et al (1990) p. 226 - 227.*

From the results of the HEGY test as presented in the table above we confirm that the seasonality in the data set can be satisfactorily modelled using seasonal dummies and further that, though not unanimously, the hypotheses of the existence of unit roots at the semi-annual and annual frequencies have been rejected at the 5% significance level. Hence we infer that the variables seem to be characterised by stochastic non-stationarities that can be removed through first order differencing. Having analyzed the stationarity characteristics of the data-set we proceed to test for the existence of cointegration among the variables in the data-set using the so-called *Johansen Procedure*.

4.3.3 Cointegration Tests

The empirical framework we have chosen in analysing the theoretical conclusions of the intertemporal approach to the current account balance requires the existence of cointegration among the variables of interest. We utilise the Johansen approach in testing for the number of cointegration vectors in the data-set. A set of variables, X_t is said to be cointegrated of order (d,b) - denoted $CI(d,b)$ - if X_t is integrated of order d and there exists a vector β, such that $\beta \mathbf{X}_t$ is integrated of order $(d - b)$[4]. The most common test for cointegration is the Engle and Granger(1987) two-step procedure which performs the tests in a univariate setup. Recent developments in the literature include the Johansen procedure (see Johansen(1988) and Johansen and Juselius(1990))[5].

[4] A variable is said to be integrated of order z - denoted $I(z)$ if the said variable becomes covariance stationary after differencing z times. See Cryer(1986) for a further definition of the concept of stationarity.

[5] These studies examine the question of cointegration and provide not only an estimation methodology but also explicit procedures for testing for the number of cointegrating vectors as well as for restrictions suggested by economic theory - in a multivariate setting.

Following Johansen and Juselius(1990) we obtain the results as summarised in Table 4.3 below. Prior to the implementation of the Johansen procedure we need to establish an appropriate/optimal lag length/order, p, for the underlying VARX. To obtain p we use multivariate lag order tests - the Akaike Information Criterion(AIC), the Swartz Criterion(SIC) and the Hannan and Quinn Criterion (HQ) - the results of which are presented in Table 4.2 below. The information criteria, as usual, are not unanimous as to the optimal lag length[6]. However guided by the statistical performance of the residuals obtained through the Johansen procedure (see Table 4.2 below[7]) we are convinced that an optimum lag of 2 is suitable for Germany and p=5 adequately captures the dynamic structure of the data-set of the US.

Table 4.2: Tests for Optimal Lag Lengths

	Tests	Number of lags							
		1	2	3	4	5	6	7	8
Germany	AIC	20.94	20.34	20.82	19.75	20.36	19.91	20.58	19.59*
	SIC	22.97*	24.41	25.92	26.87	28.52	30.09	31.79	32.85
	HQ	21.64	21.73	21.91	21.52*	21.83	22.08	22.44	22.15
US	AIC	19.01	19.38	18.85	19.25	18.77*	19.49	18.94	19.26
	SIC	21.14*	22.64	24.24	25.77	27.42	29.27	30.85	32.31
	HQ	19.74*	19.84	20.05	20.18	20.42	20.88	21.06	21.12

Notes: *All statistics are calculated using multivariate methods. The starred numbers in each row indicate the minimum value attained (and hence correspondingly the optimum lag selected) by the respective information criteria.*

[6] Paulsen(1984) has shown that both the Swartz Bayesian criterion (SIC) and the law-of-iterated-logarithm Criterion (LIL) of Hanna and Quinn are weakly consistent in the presence of unit roots whereas the Akaike Information Criterion (AIC) asymptotically over-estimates the optimal lag length lag.

[7] The statistical performance of residuals obtained using other lag lengths not only yield insufficient evidence for the rejection of the null hypothesis of no cointegration vectors (in some cases) but also leave much to be desired. In estimating the cointegrating vectors we allow for trend in the data.

For p=2 for Germany (and p=5 for the US) for we perform the *trace* and *lambda max* tests using the Johansen Procedure. From the results below we infer that it is reasonable, using the *trace* test, to accept the null of two cointegrating vectors at the 10% significance level for both countries. Univariate residual diagnostic tests are performed using the Ljung-Box test for autocorrelated residuals, ARCH test for autoregression and conditional heteroscedasticity, the skewness and excess-kurtosis statistics, and the Jacque-Bera test for normality. The test results as depicted in table 4.4 below reveal no indications of misspecification error in the estimated VARX model based on the existence of two cointegrating vectors. However the likelihood ratio test results indicate that investment and the current account balance series are stationary at the 5% significance level - a finding that seems contrary to the results of the seasonal integration results presented in table 4.1 above. Considering that the former test deals mainly with the *residuals* of the said variable (within a multivariate context) whereas the HEGY tests - just like all the other tests for unit roots - tests the *variable itself* directly we proceed with the paper based on the results of the HEGY tests.

Table 4.3: Initial Results of Cointegration Analyses

A.The Maximum Eigenvalue (λ_{max}) Criteria.

Hypothesis		Test Statistics		Critical Values	
H_0	H_1	Germany	US	90%	95%
r=0	r=1	28.67	26.14	24.73	27.07
r≤1	r=2	20.14	15.48	18.60	20.97
r≤2	r=3	6.43	9.91	12.07	14.07
r≤3	r=4	2.80	3.51	2.69	3.76

B. The Trace Criteria.

Hypothesis		Test Statistics		Critical Values	
H_0	H_1	Germany	US	90%	95%
r=0	r≤1	58.05	55.04	43.95	47.21
r≤1	r≤2	29.38	28.90	26.79	29.68
r≤2	r≤3	9.23	13.32	13.33	15.41
r≤3	r=4	2.80	3.51	2.69	3.76

Notes: *An intercept and three seasonal dummy variables are included in each estimated equation. The reported critical values are taken from Osterwald-Lenum (1990).*

c. Likelihood Ratio (LR) Tests for Stationarity (Given r=2).

Null Hypothesis	$y_t \sim I(0)$	$c_t \sim I(0)$	$inv_t \sim I(0)$	$ca_t \sim I(0)$
Germany	16.009 (0.0009)	14.169 (0.0008)	5.854 (0.054)	4.301 (0.116)
USA	6.509 (0.039)	6.308 (0.043)	5.417 (0.067)	5.672 (0.059)

Notes: *The test statistics here have asymptotic $\chi(k)$ distribution where k is the number of common trends. the numbers in parentheses are estimated significance levels.*

d. Estimated Cointegration Vectors[8]

$$\beta^{Germany} = \begin{bmatrix} 0.07 & -0.17 & 0.33 & 1 \\ -0.09 & 0.16 & 0.02 & 1 \end{bmatrix} \qquad \beta^{US} = \begin{bmatrix} -0.25 & 0.27 & 0.62 & 1 \\ -0.54 & 0.71 & 0.40 & 1 \end{bmatrix}$$

[8] We do not try to identify and interpret the cointegrating relationships here since we are not interested in these equilibrium relationships as such. Note that the estimates of the Common Trends model are not sensitive to the mode of normalisation of the estimated cointegration vectors.

Table 4.4: Univariate Residual Analysis

A. Germany (Given r = 2)

	Residual Diagnostics				
Equation	L-B(17)	ARCH(2)	Skew.	Ex-Kurt.	J-B.Norm.
Δy	29.186	0.075	-0.619	1.328	9.605
ΔC	24.331	0.838	-0.074	0.449	0.651
Δinv	16.369	2.753	-0.703	0.568	6.699
Δca	22.052	0.114	1.508	6.291	141.965

B. US (Given r = 2)

	Residual Diagnostics				
Equation	L-B(16)	ARCH(5)	Skew.	Ex-Kurt.	J-B.Norm.
Δy	22.648	4.687	0.034	0.710	1.419
ΔC	16.212	3.702	-0.225	0.232	0.713
Δinv	12.075	1.309	-0.090	0.336	0.404
Δca	6.631	3.572	1.083	2.953	37.442

Notes: *The entries under the Ljung-Box (L-B(17)) are test statistics for autocorrelation and have $\chi^2(16)$ distribution. Under the column labelled ARCH(1) are statistics for testing autoregression and conditional heteroschedsticity and have $\chi^2(1)$ distributions. The next two columns are statistics for testing skewness and excess kurtosis respectively. They have $\chi^2(1)$ distributions. The next column entries are statistics for the Jarque-Bera test for normality (with a $\chi^2(2)$ distribution) based on the Skewness and Excess Kurtosis statistics.*

4.4 The Estimated Common Trends Model

From the preliminary data analyses and the results from implementing the Johansen procedure as presented for both countries in the previous sections we infer that i). the hypothesis of stationarity in levels is strongly rejected for all of the series and hence the data-set is characterised by stochastic non-stationarities that can be removed by first-difference transformations; and ii). the data-set is characterised by two cointegrating vectors as represented table 4.3 above - implying the long run peculiarities of the data-set are driven by two common stochastic trends. It seems therefore most appropriate to conclude that an error-correction model with two cointegrating vectors is a reasonably congruent representation of the data-set in the cases of both countries. Deducing from the theoretical framework presented in section 4.2 above we postulate that the productivity shocks contain both country-specific and global components. We indicate these shocks as τ_{Dt} and τ_{Gt} respectively. The exact identification (and estimation) of these trends requires some restriction(s) on the A matrix of the common trends model as described in chapter 1 of this dissertation and reproduced in the equation below.

$$X_t = X_0 + A\tau_t + C^*(L)e_t \qquad (4.18)$$

where $\tau_t = \mu + \tau_{t-1} + \varphi_t$ and $X_t = [\ y_t,\ c_t,\ inv_t,\ ca_t\ ]'$. Since we have $k\ (= n - r) = 2$ common trends - where r indicates the number of cointegration vectors in the n-variable data-set - the A matrix above must necessarily contain $n{\times}k$ (= 8) elements. From the requirements already derived [see chapter 1 of this dissertation] as $\beta'A = 0$ (yielding r×k restrictions/equations) and $C(1)\Xi C(1)' = AA'$ (which yields $k(k+1)/2$ restrictions) we obtain a total of $rk + k(k+1)/2$ (= 7) equations. Hence to exactly identify the A matrix we require only one additional restriction. The most suitable candidate in this case (and that is also consistent with the theoretical model presented/discussed above is that which states that global shocks, because they affect all countries, have no permanent effects on the current accounts of these countries[9]. To choose the structure of the A matrix to suit/effect this particular restriction note that for estimation purposes $A = A_0\Pi$ where A_0 is an $n{\times}k$ initial impact matrix and Π is $k{\times}k$ lower triangular. We find, in this particular paper, the choice of the initial impact matrix, A_0, of the form

[9] Notice here that given that we have two stochastic trends in the data-set for each country, the imposition of this theoretical a priori restriction implies the imposition of an additional necessary identifying restriction.

$$A_0 = \begin{bmatrix} 0 & 1 \\ 1 & \alpha_{22} \\ \alpha_{31} & \alpha_{32} \\ \alpha_{41} & 0 \end{bmatrix}$$

very suitable since it produces the desired structure of A $(= A_0\Pi)$ with the embedded proposition of no long-run effect of global productivity shocks on the current account balance given that $\tau_t = [\tau_{Dt}\ \tau_{Gt}]'$ [10].

Having discussed the necessary restrictions for distinguishing between country-specific and global shocks we turn next to the quantitative effects that the data assigns each of these shocks. Using a **VAR(p)** with p=2 in the case of Germany and p=5 in the case of the US the estimates of the **CT** model are as follows.

The asymptotic standard errors in parentheses (where * indicates statistical significance at the 5% level) are obtained under the assumption of normality (see Theorem 2.3 on page 23-25 of Warne(1990)). The coefficients measure the long-run effects of the respective stochastic trends on the corresponding elements of the X_t-vector. In what follows we discuss the empirical results seeking answers to the following specific questions asked in the introduction: i) what are the long run effects of each of these shocks (analyzed separately) on investment and the current account? ii) how does each of our variables of interest (investment and the current account) respond to one standard deviation innovations in the respective shocks? iii) what is the relative performance of the innovations in the respective shocks in

[10] Given the interrelatedness of shocks across countries the most appropriate way to conceptualise these shocks is to think of them as the outcome of a number of unidentified system-wide shocks (having effects across country borders) rather than as a set of country-specific shocks. However, as we do here, the imposition of restrictions/implications derived a priori from economic theory may provide acceptably convincing means of identifying the effects of a specific type of shocks to any particular country under consideration.

$$\text{Germany:}\quad \begin{bmatrix} y_t \\ C_t \\ Inv_t \\ Ca_t \end{bmatrix} = X_0 + \begin{bmatrix} 6.317 & 30.047^* \\ (12.895) & (10.906) \\ 7.391 & 16.526^* \\ (7.493) & (5.998) \\ 4.594^* & 2.103^* \\ (1.471) & (0.763) \\ -0.705^* & 0 \\ (0.135) & (-) \end{bmatrix} \begin{bmatrix} \tau_{Dt} \\ \tau_{Gt} \end{bmatrix} + C^*(L)e_t$$

$$\text{US:}\quad \begin{bmatrix} y_t \\ C_t \\ Inv_t \\ Ca_t \end{bmatrix} = X_0 + \begin{bmatrix} -4.032 & 43.006^* \\ (2.610) & (16.174) \\ 0.869 & 30.534^* \\ (0.871) & (11.484) \\ 7.059 & 4.309^* \\ (4.547) & (1.617) \\ -5.635^* & 0 \\ (2.713) & (-) \end{bmatrix} \begin{bmatrix} \tau_{Dt} \\ \tau_{Gt} \end{bmatrix} + C^*(L)e_t \qquad (4.19)$$

explaining fluctuations in investment and the current account? and finally iv) how does our approach compare with that of G&R? The first question is attempted using the results from the estimated common trends model. More specifically the answer to this question is based on the estimates of the A matrix as reported for both countries in (4.19) above. Forecast error variance decompositions, as reported in table 4.5, aid us in tackling the third question whereas the second question is addressed using estimated impulse response functions as reported in figures 4.2 through 4.5 for both countries. The conclusion provides some answers to the final question.

In conformity with the predictions of the theory the current account balance is significantly affected by permanent domestic(country-specific) shocks. Thus whereas a positive domestic (country-specific) shock significantly worsens the current account of the US by as much as 5.64 billion US dollars the corresponding figure in the case of Germany is 0.71 billion deutsche marks as shown by the estimates of the A matrix in (4.19) above. Permanent global shocks (by restriction derived from the theory and imposed as an identifying assumption) do not have any long run effects whatsoever on the current account balance. This result is depicted by the zero coefficient in the estimated A matrix for each country. These shocks have however very significant positive long run effects on investment in both countries. More specifically, whereas permanent global shocks significantly increase investment by as much as 4.31 billion US dollars in the US the corresponding effect of these shocks in Germany is an estimated increase of 2.10 billion deutsche marks in investment. These results are very consistent with our theoretical predictions as outlined in section 4.2. Permanent domestic shocks do have positive effects on investment in both countries.

Plots of the impulse response functions of each of the elements of $\boldsymbol{X}_t = [y_t\ c_t\ inv_t\ ca_t]'$ to a one standard deviation innovation in the permanent stochastic trends and their transitory counterparts are reported in figures 4.2, 4.3, 4.4 and 4.5 where the responses over 40 quarters are plotted with their 95% confidence bounds. As reported in figure 4.2 responses of these variables to the permanent domestic shocks are not significant at any of the horizons even though there is a clear indication of the direction of the effects especially as regards investment and the current account in both countries. Investment responds positively to these shocks whereas the response of the current account is negative over the entire horizon of the impulse response functions in both countries. Permanent innovations in the global trend however lead generally to significantly positive responses in output, consumption and investment in the two countries over the entire horizon of the estimated impulse response functions. The current account, by restriction as imposed on the A matrix before estimation, does not respond to shocks to these permanent global shocks except during the first three quarters in Germany when the effect is positive but moves rapidly to zero. We turn now to discussing the effects of the transitory shocks as reported in figures 4.4 and 4.5. Except for their effects on investment during the first five quarters (following a one standard deviation shock) in the case of Germany transitory (domestic and global) shocks do not have any significant influences on the endogenous variables. Even in the case of Germany the effects of innovations in transitory domestic and global shocks on consumption and investment respectively are but *transitory* and move rapidly to null within five quarters. This indicates that the productivity shocks - domestic/country-specific and global - can be considered as following a random walk processes (ie. the absolute value of the autoregressive parameter in equation (4.2) on page 76 is not significantly different from one - i.e. $|\rho| \approx 1$)[11]. This result is confirmed by the forecast error variance decompositions (the results of which are reported in table 4.5 below) showing that all fluctuations in the respective variables are exclusively explained by permanent shocks.

Having discussed the dynamic responses of each variable with respect to innovations in the two stochastic trends we move on next to examine the relative significance of these innovations in explaining long run fluctuations in investments and the current account balance. To do this we present the results of our forecast error variance decompositions showing the importance of each shock in explaining the variability of the forecast error in respect of each variable in the long run[12]. Permanent shocks to domestic (country-specific) trends explain between 73% and 83% of total long run fluctuations in investment and are solely responsible (by restriction as required by the theory) for all variations in the current account.

[11] This result showed up in G&R as well. Using standard univariate unit root tests G&R could not reject this $\rho = 1$ hypothesis for any of the countries considered. Hence all their subsequent empirical results assumed $\rho = 1$ - that is to say the shocks are exclusively permanent.

[12] Forecast error variance decompositions for different time horizons are not presented here since they indicate proportions of variations in the respective variables within the specified short run horizons (horizons on which our theoretical model gives very little information, if any) explained by innovations in the respective trends.

Table 4.5: Long-Run Forecast Error Variance Decompositions

	Domestic Shocks (τ_{Dt})		Global Shocks (τ_{Gt})	
	Germany	US	Germany	US
Output (y_t)	0.042 (0.162)	0.009 (0.113)	0.958* (0.162)	0.991* (0.133)
Consumption (c_t)	0.167 (0.279)	0.001 (0.035)	0.833* (0.279)	0.999* (0.282)
Investment (inv_t)	0.827* (0.129)	0.729* (0.282)	0.173 (0.129)	0.271 (0.282)
Current Acc. (ca_t)	1.000 (-)	1.000 (-)	0.000 (-)	0.000 (-)

Notes: *The estimated standard errors are based on Theorem 2.4 of Warne(1990). Very insignificant or zero standard errors are indicated by (-) in the table. Again, a * indicates statistical significance at the 5% level.*

Global shocks explain 17% (27%) of total long run variations in investment in Germany (the US) and domestic shocks are accountable for 83% (73%) of long run variations in the same variable in the respective countries. Further, variations in output and consumption are explained solely by global shocks - domestic shocks explain rather low and statistically insignificant proportions of variations in both variables. More specifically, given our identifying assumption (that global shocks have no long run effects on the current account) global shocks tend to explain the bulk of fluctuations in income and consumption whereas domestic shocks explain between 70% and 80% of fluctuations in investment. What is the likely explanation for this? It may be due to the fact that the adjustment costs associated with investment have their origins in domestic structural/institutional arrangements that are more responsive to domestic shocks than they are to global shocks.

4.5 Summary and Conclusion

Using the common trends approach this paper identifies and analyses the relative effects of idiosyncratic and global stochastic shocks that influence investment and the current account.

To be able to identify these stochastic shocks we use the intertemporal approach to the current account in the tradition of G&R, Obstfeld(1986) and Razin(1993). The intertemporal approach acts as the background theory for explicit specification and analyses of the two shocks - domestic (country-specific) and global - and provides the long run restriction(s) that is(are) used not only in identifying crucial matrices during the estimation of the common trends model but also in explaining the results of our impulse response experiments and variance decompositions.

Our empirical estimates yield the following stylized facts. Generally the long run effects of domestic (country-specific) productivity shocks on the current account are significantly negative as the theoretical model predicts. The estimates of the A matrix of our common trends model however reveal that investment responds positively to both permanent innovations; responding more to domestic (country-specific) shocks than to global shocks. Nevertheless the apparent puzzle that the current account responds, in absolute terms, by much less than investment to domestic (country-specific) productivity shocks - as pointed out by G&R - seems to persist even with the approach adopted here. In our view the solution to this apparent puzzle requires a research strategy that explicitly models, estimates the parameters of the cost-of-adjustment investment technology and examines relative responses of investment and the current account to the two productivity shocks for given parameter values of the cost of adjustment of investment. Despite this the common trends technique used in this paper adequately describes fluctuations in the data-set and yields results that are consistent not only with the intertemporal model of the current account as adopted here but also with most aspects of earlier empirical work on the subject. Hence it is not unreasonable to conclude that the common trends approach is a potential alternative to estimations based on calculated Solow residuals and gives the researcher more information on the dynamics of the effect of innovations in productivity shocks - a dynamics that is very crucial in empirical investigations on capital mobility and the related issues of current account fluctuations.

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Figure 4.1a: Seasonal Analyses of German GDP, y_t

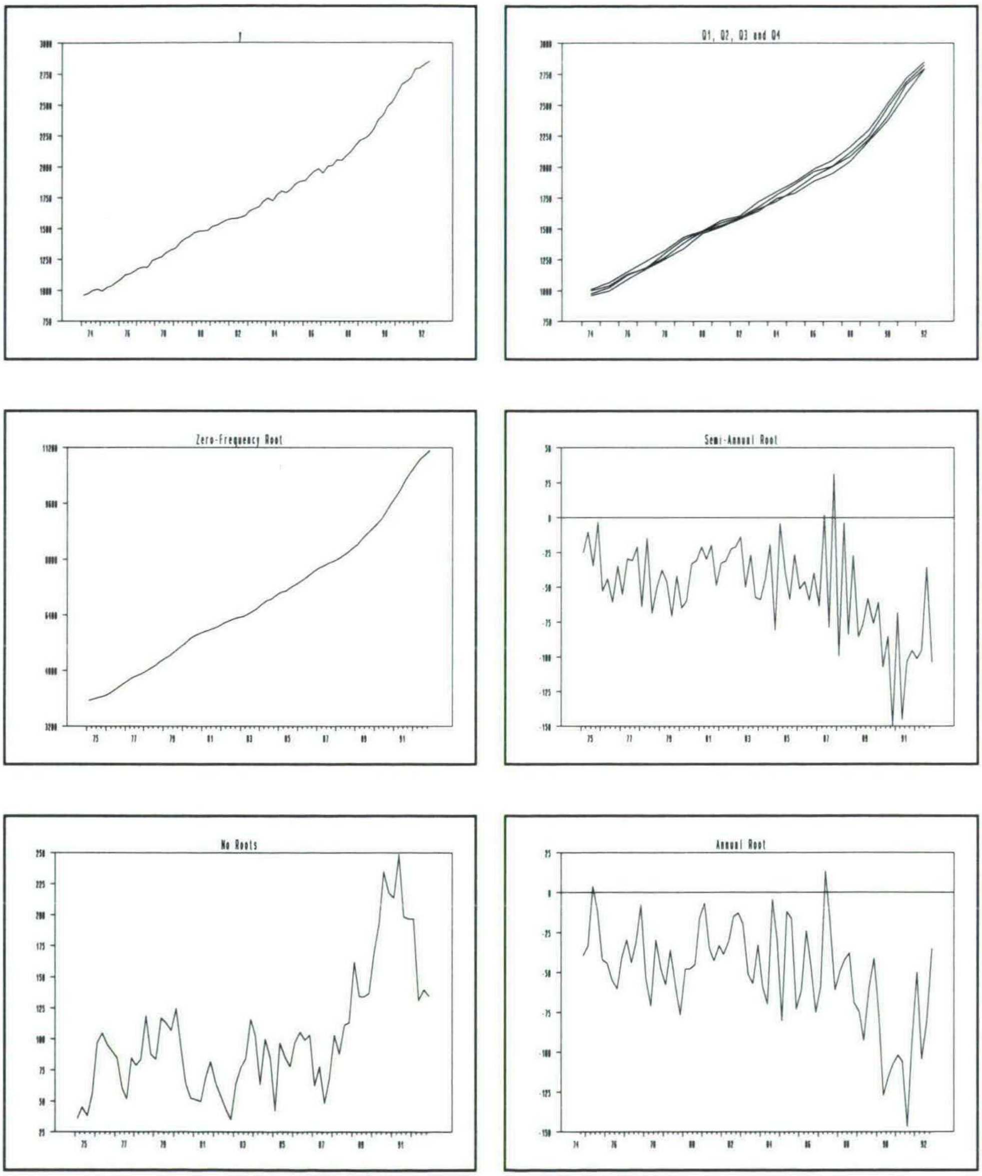


Figure 4.1a (contd.): Seasonal Analyses of German Private Consumption, c_t

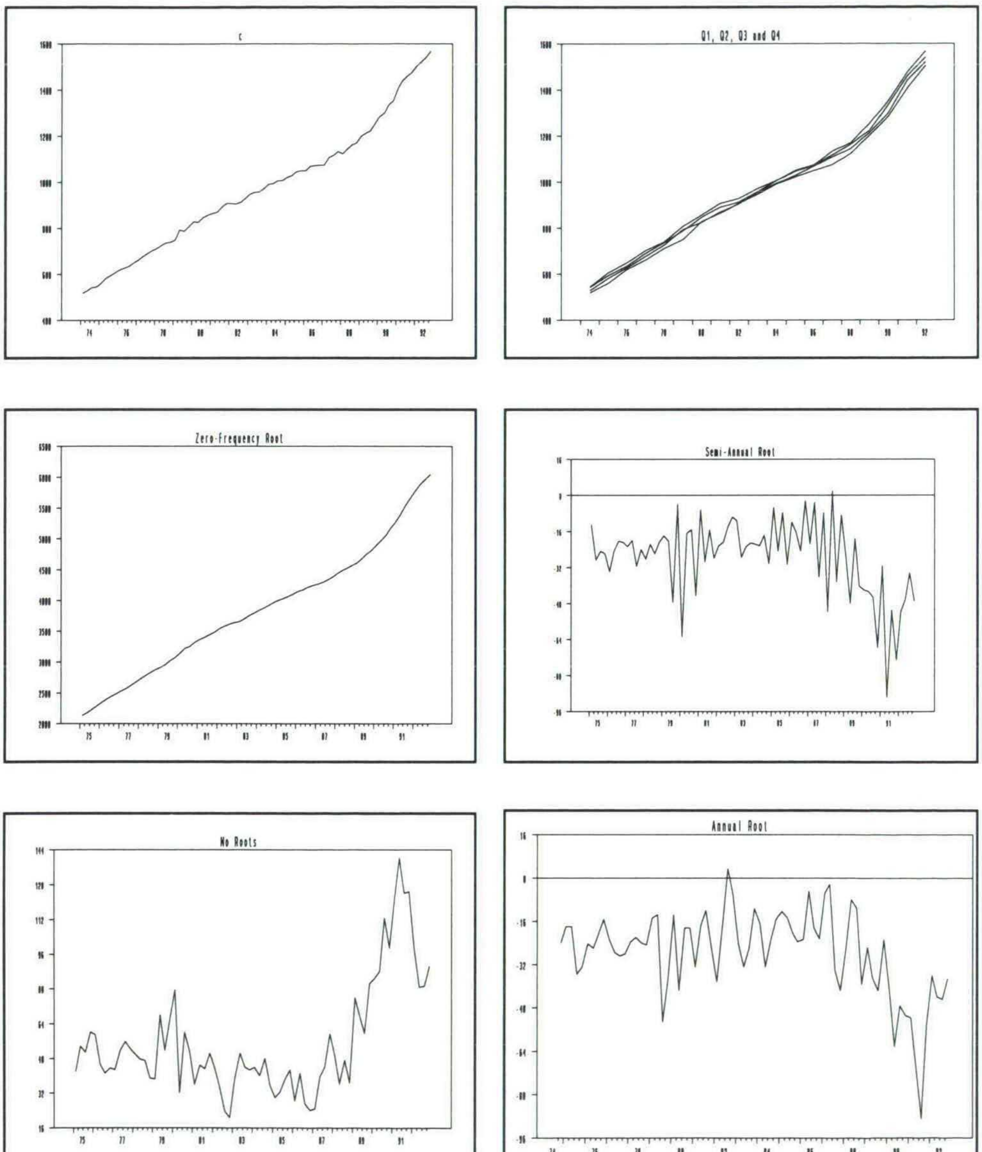


Figure 4.1a (contd.): Seasonal Analyses of German Gross Investment, inv_t

Figure 4.1a (contd.): Seasonal Analyses of German Current Account Balance, ca_t

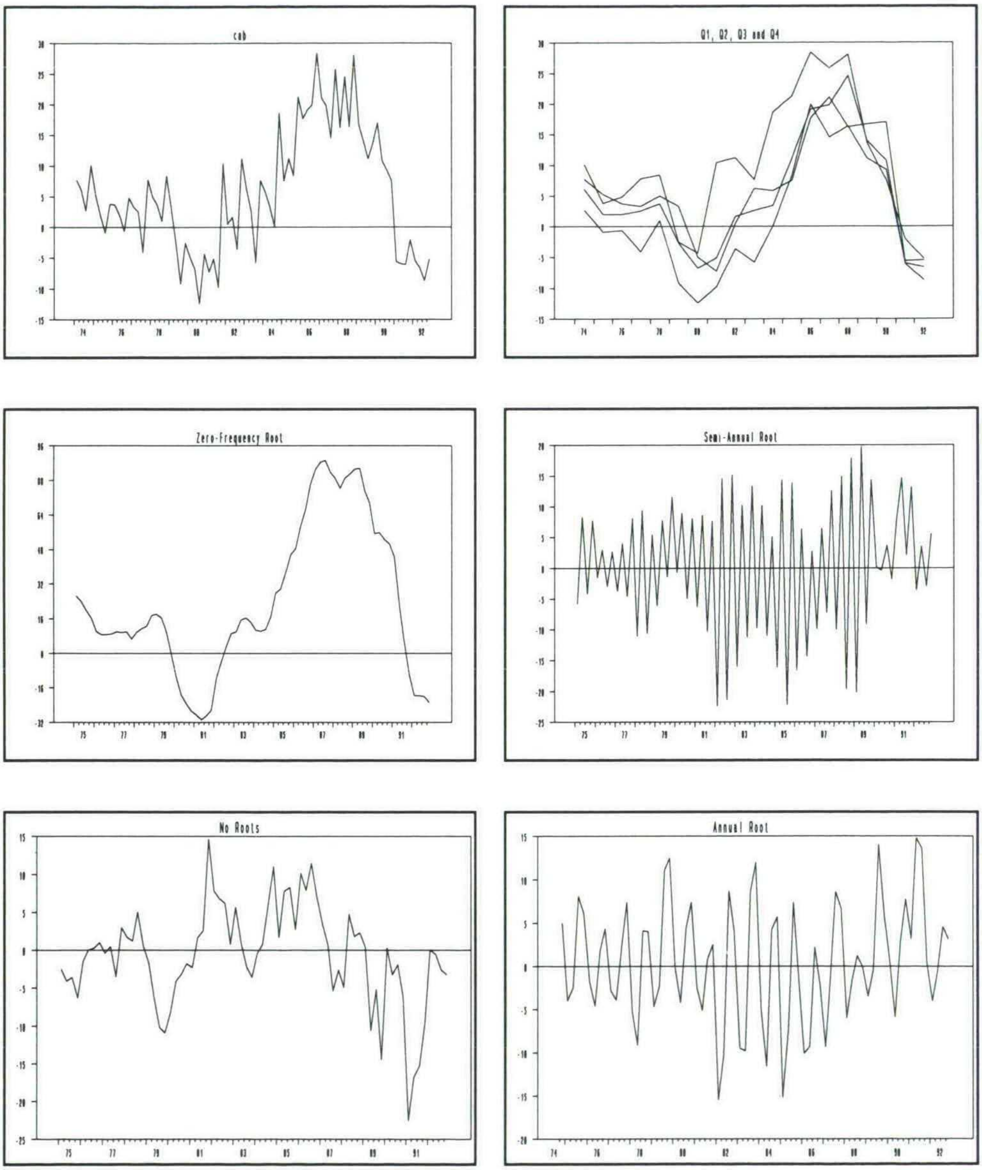


Figure 4.1b: Seasonal Analyses of US GDP, y_t

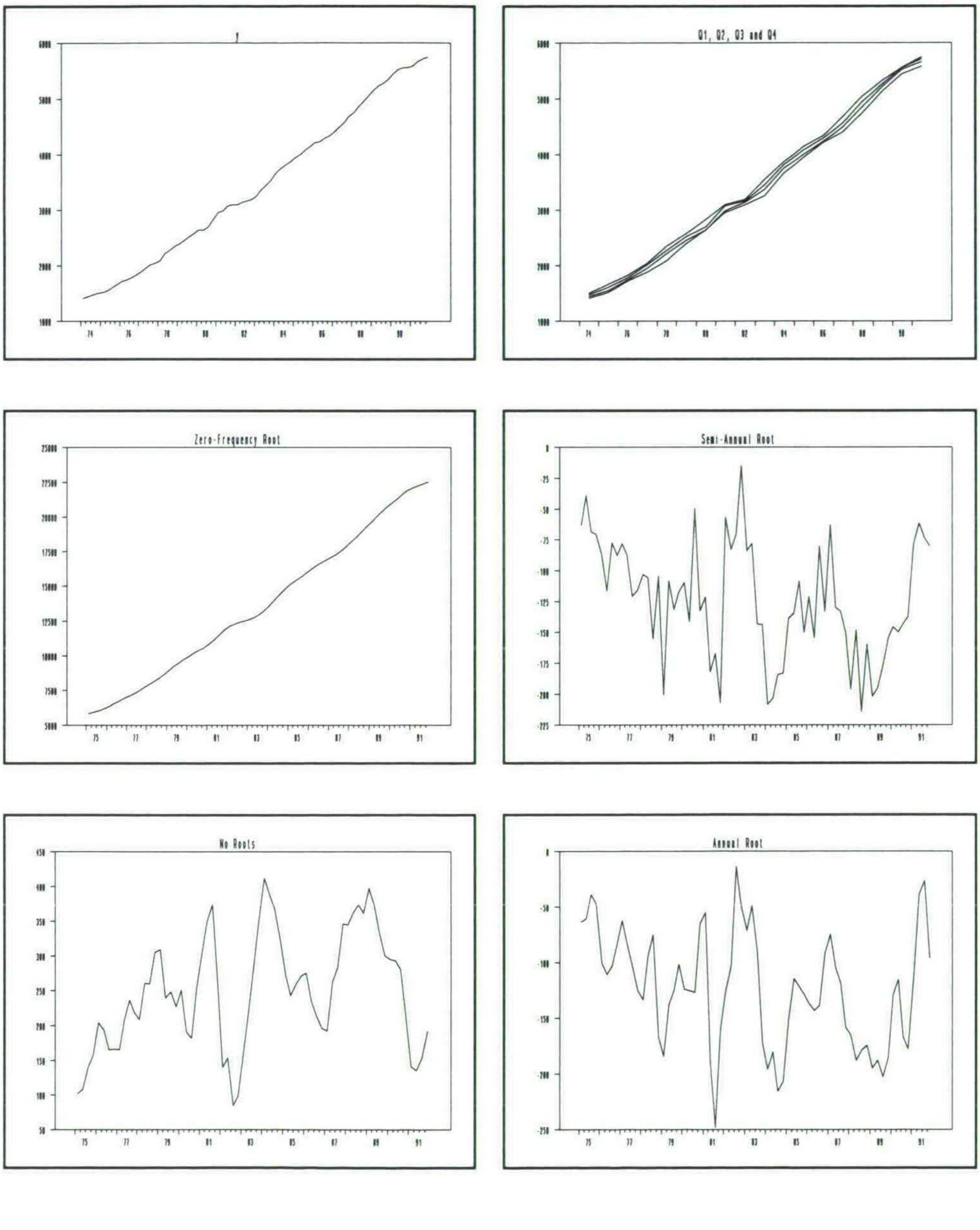


Figure 4.1b (Contd.): Seasonal Analyses of US Private Consumption, c_t

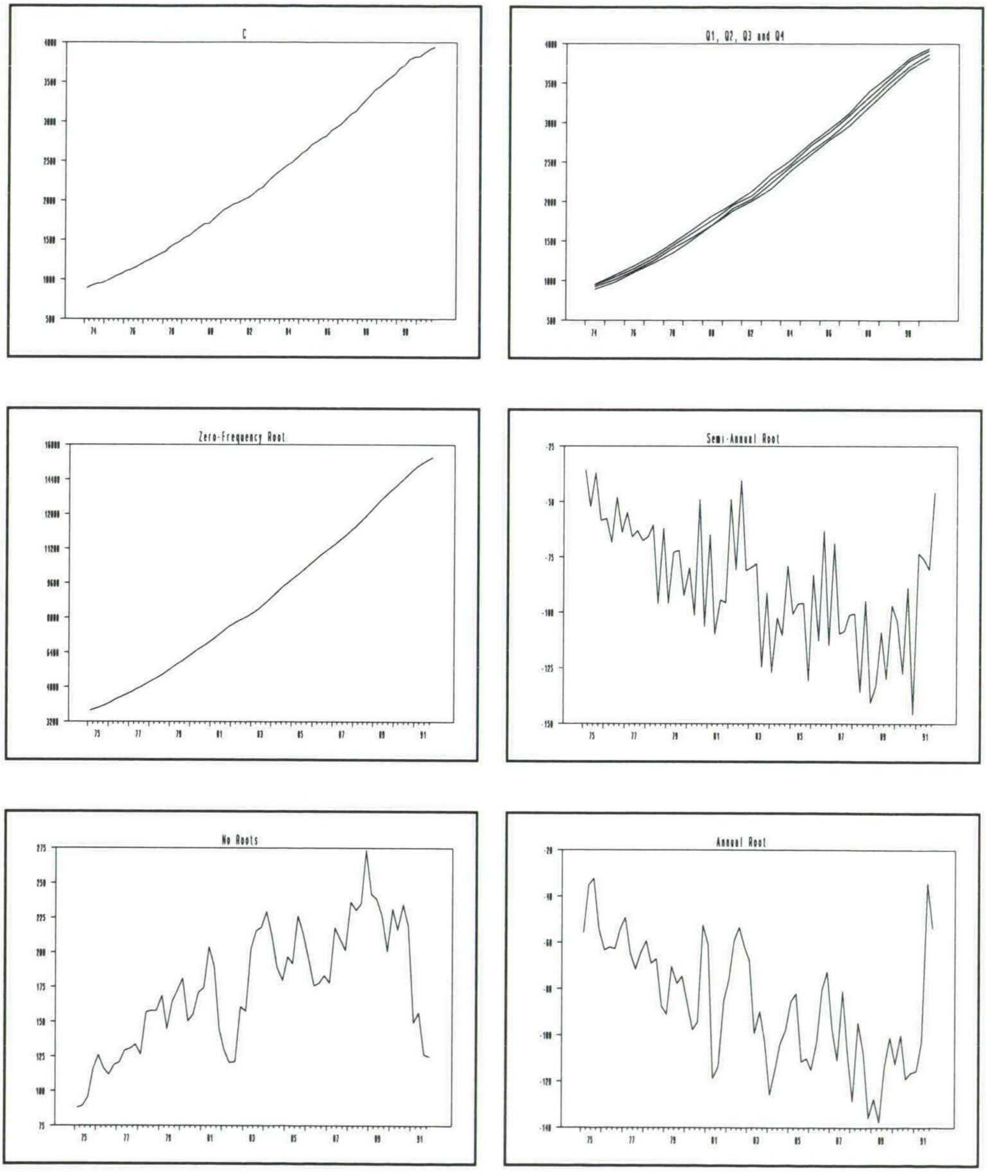


Figure 4.1b (Contd.): **Seasonal Analyses of US Gross Investment, inv_t**

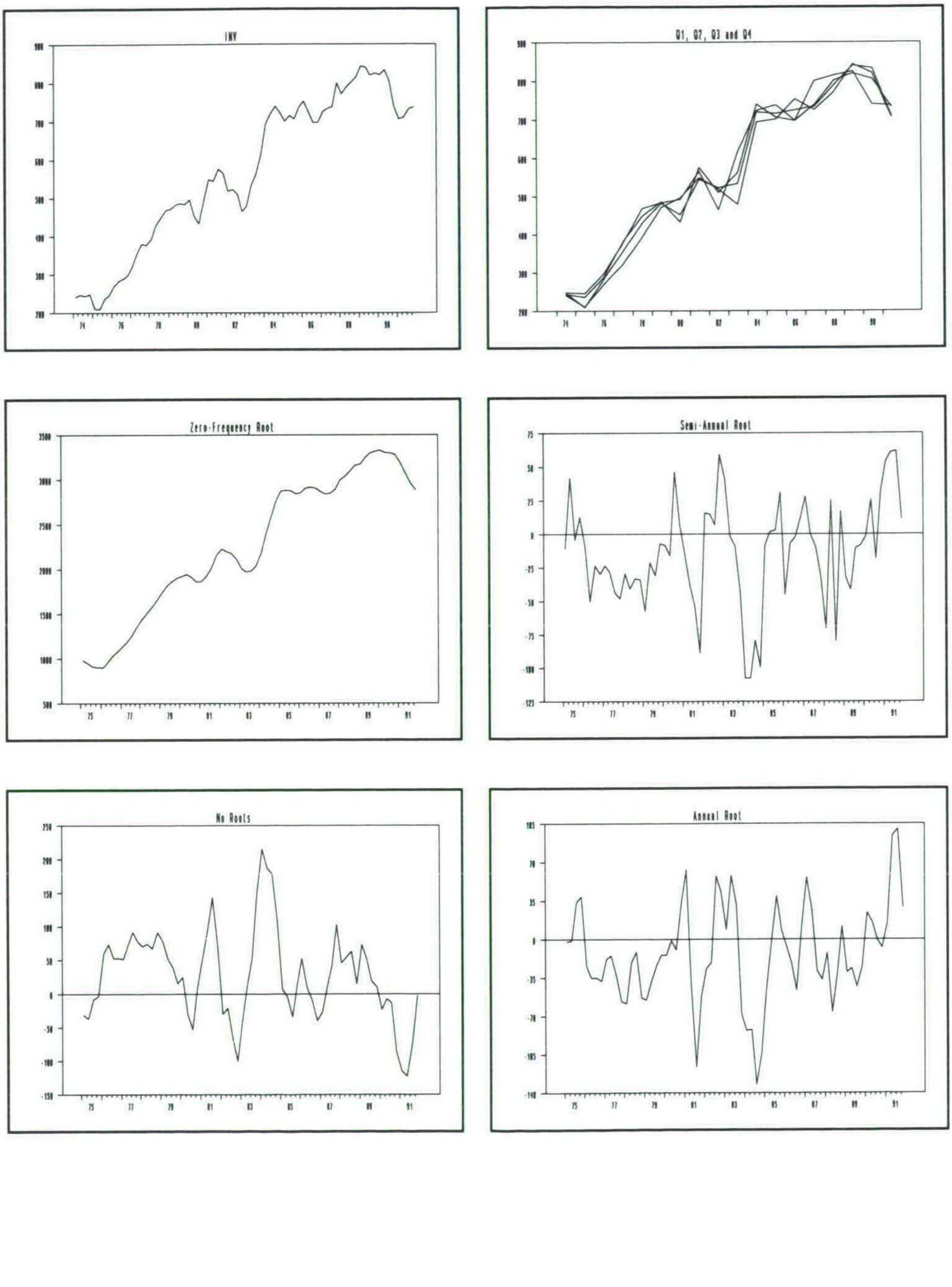


Figure 4.1b (Contd.): Seasonal Analyses of US Current Account Balance, ca_t.

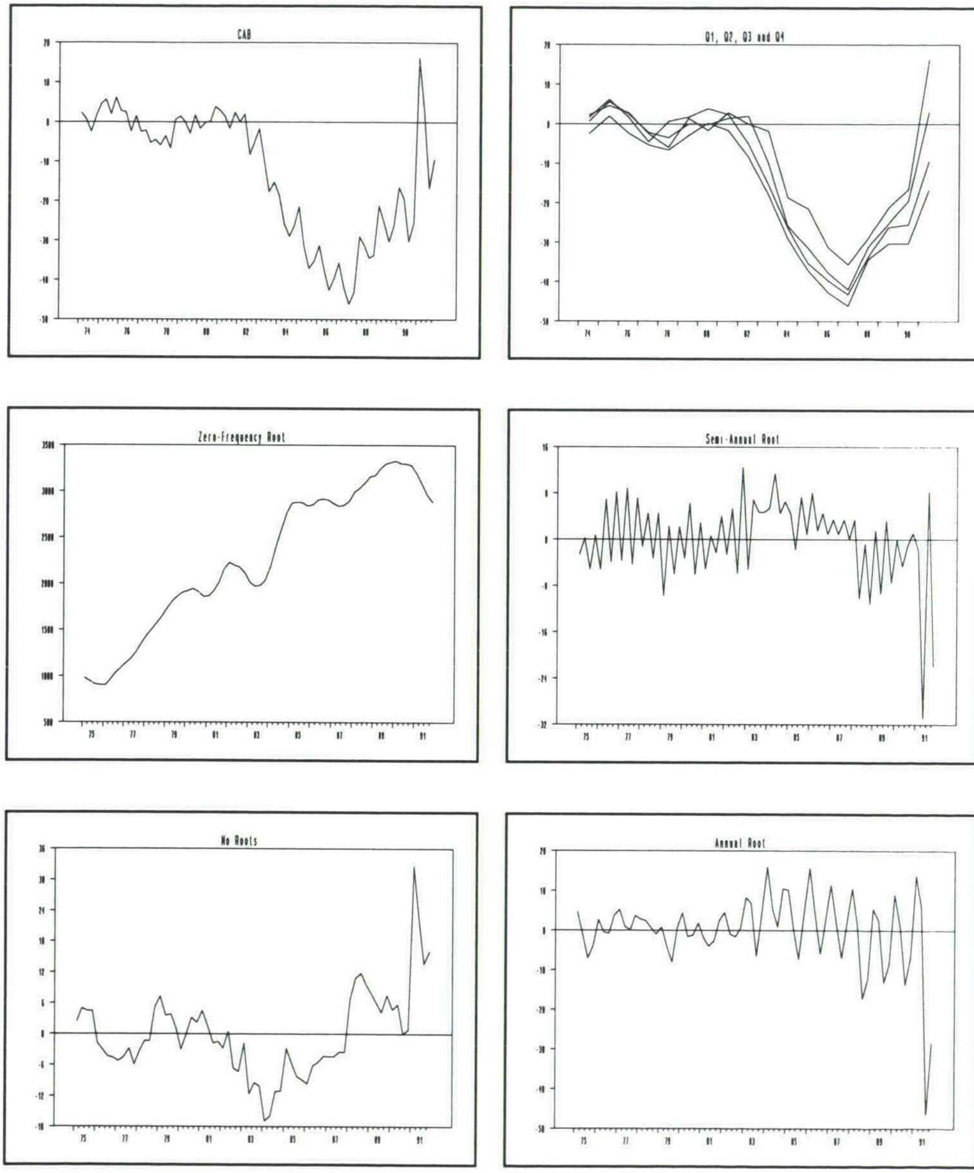


Figure 4.2: Impulse response functions (with 95% Confidence Intervals) from a one standard deviation shock to the permanent domestic (country-specific) productivity trend (τ_{Dt})

A. Germany

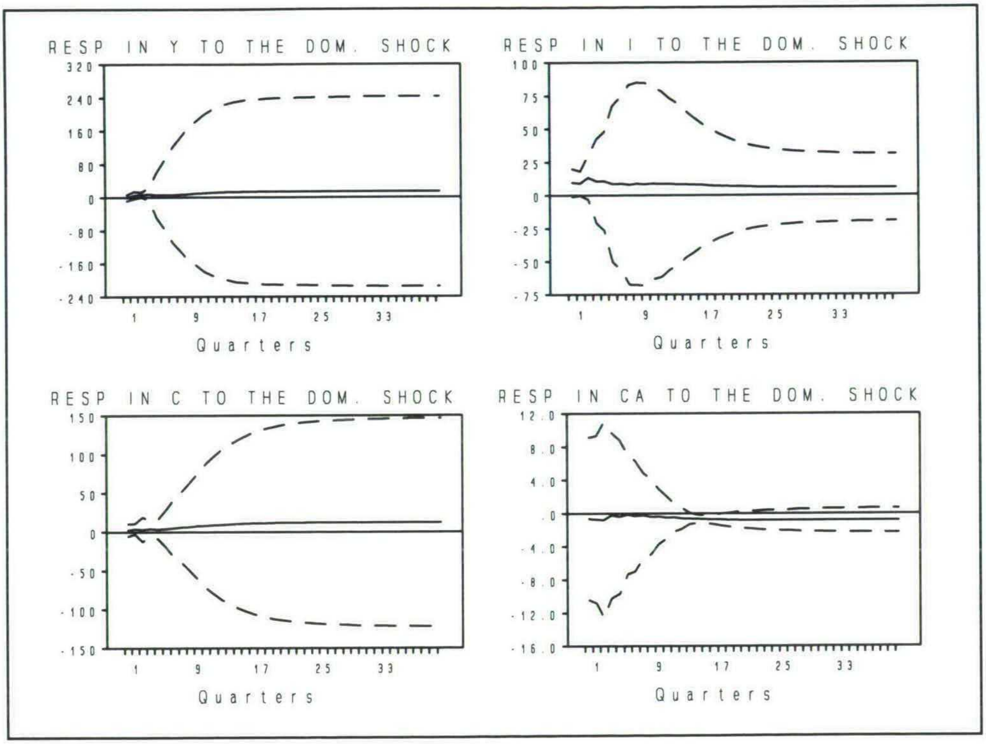


Figure 4.2 (contd.):

B. US.

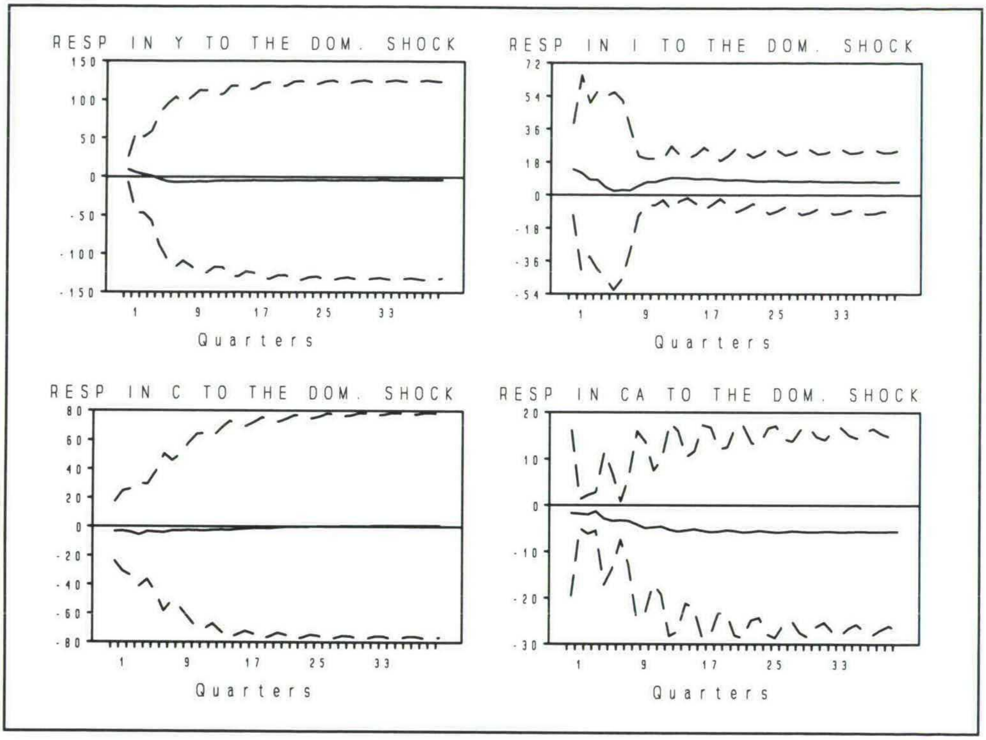


Figure 4.3: Impulse response functions (with 95% Confidence Intervals) from a one standard deviation shock to the permanent global trend (τ_{Gt})

A. <u>Germany</u>

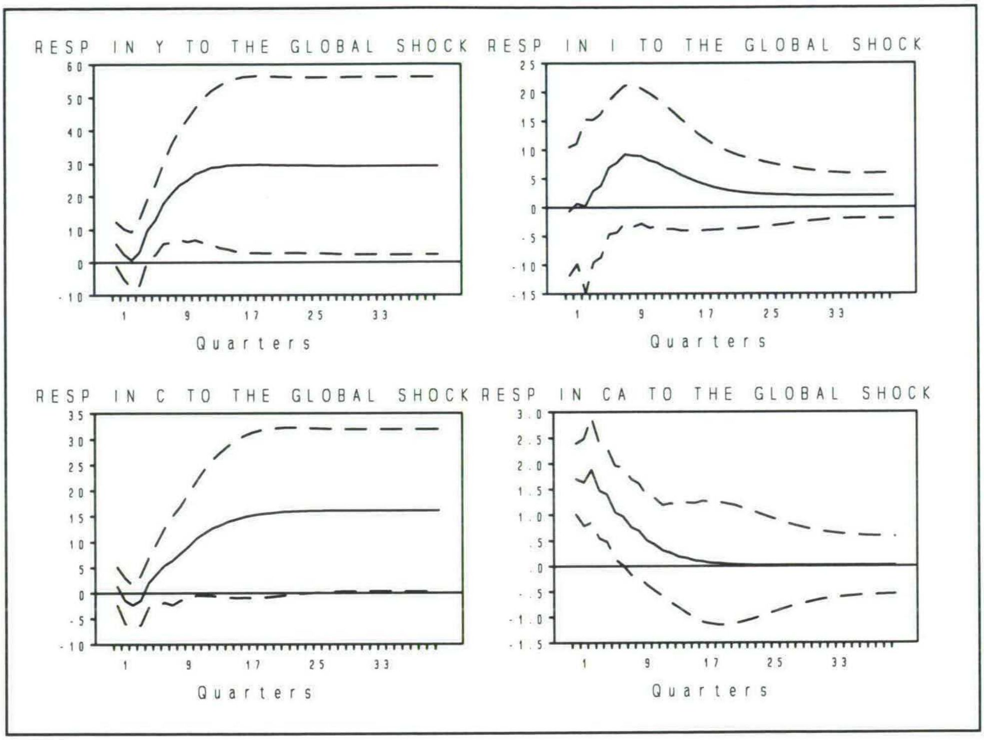


Figure 4.3 (contd.):

B. US

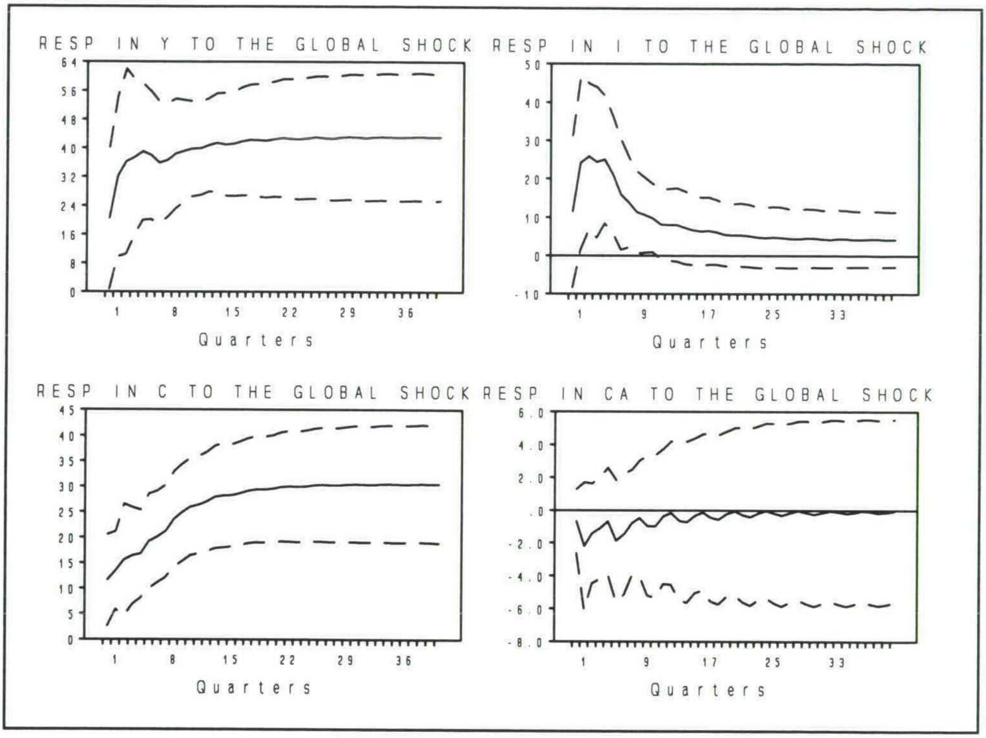


Figure 4.4: **Impulse response functions (with 95% Confidence Intervals) from a one standard deviation shock to the transitory domestic (country-specific) productivity trend**

A. Germany

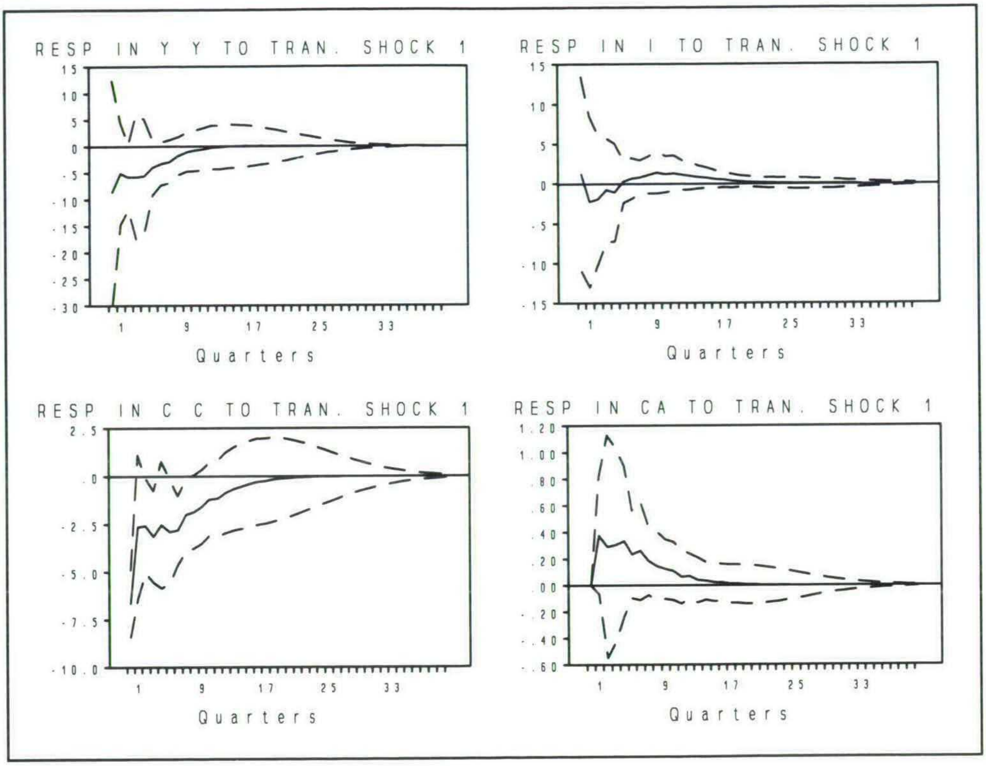


Figure 4.4 (contd.):

B. US

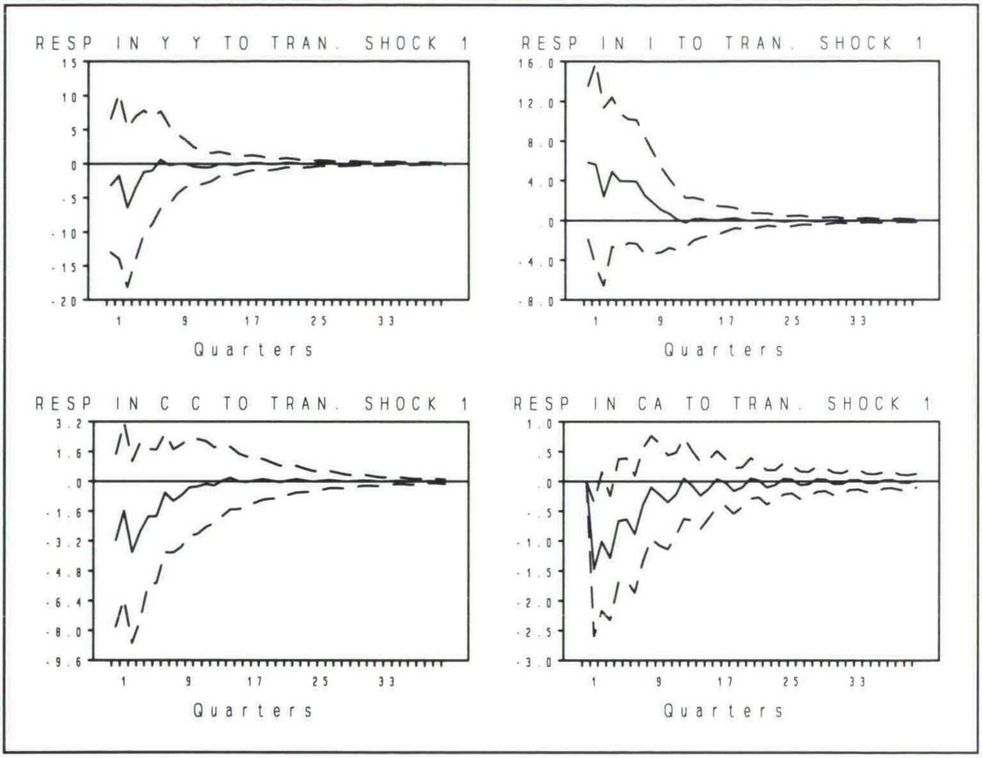


Figure 4.5: Impulse response functions (with 95% Confidence Intervals) from a one standard deviation shock to the transitory global productivity trend

A. Germany

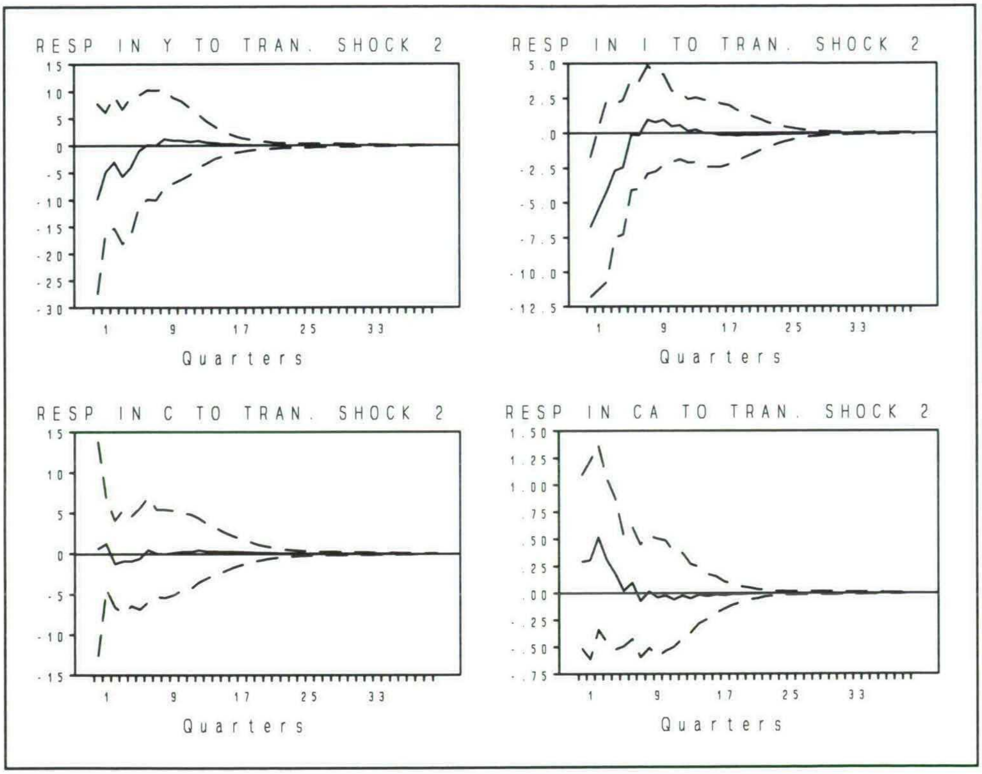
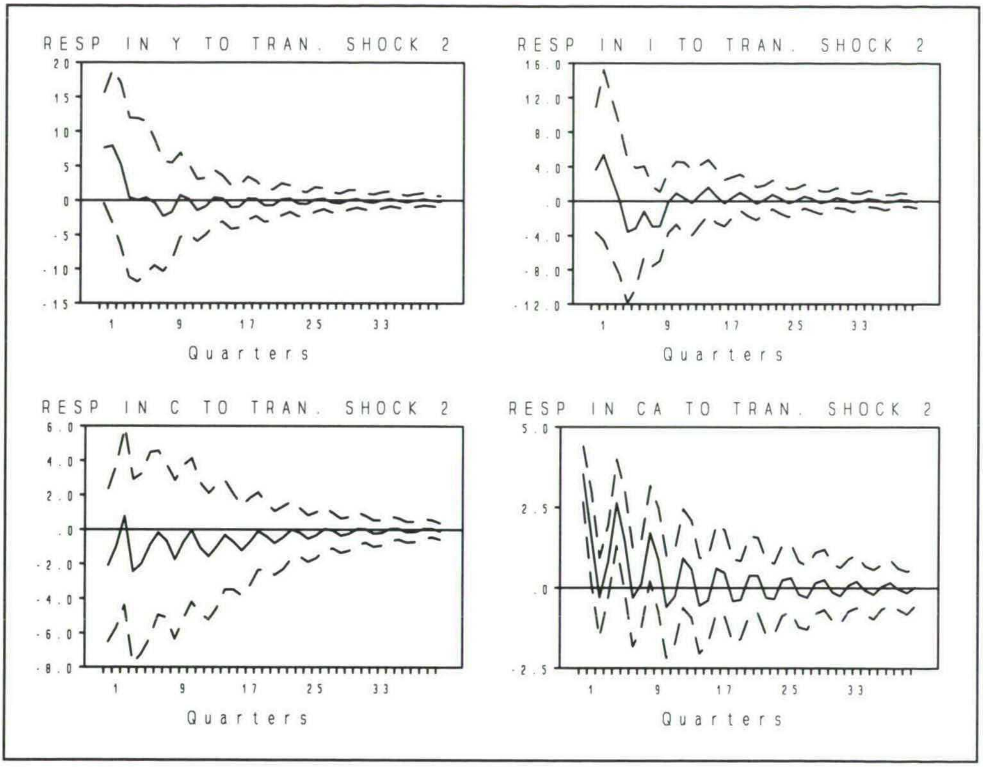


Figure 4.5 (contd.):

B. US



Chapter 5:

The Effect of Monetary Policy on Exchange Rates: How to Solve the Puzzles

5.1 Introduction

In response to the opportunity offered by the post Bretton Woods period of generalized floating and excessively volatile exchange rates a substantial literature on the linkages among macroeconomic variables and how these linkages explain observed exchange rate volatility has emerged. We shall be concerned here with only that strand of the literature that models responses of exchange rates to monetary policy shocks. In particular our reference models involve those based on asset markets. Most of the empirical results obtained from tests of these models leave much to be desired. Dornbusch(1979) and Frenkel(1984) however are able to find evidence in support of the flexible price monetary model of exchange rates even though their results are not robust to sample period extensions. MacDonald and Taylor(1993) use cointegration methods (viewing the monetary model as a long-run equilibrium condition to which the exchange rate converges) to empirically test the implications of the monetary approach to exchange rates. The portfolio balance model (PBM) has been empirically tested by Branson, Halttunen and Masson(1977) using a reduced form exchange rate equation derived from the PBM - the results revealed insignificant coefficients and the persistence of autocorrelation in the residuals. As we discuss below recent attempts at empirical analysis of the effects of monetary policy shocks on exchange rate fluctuations has encountered a number of puzzles.

We propose in this paper a new method for estimating the contemporaneous and dynamic relationships between monetary policy shocks and movements in bilateral exchange rates - nominal and real - based on the current state of the debate in the monetary policy literature on the measurement of policy shocks. Traditionally, the stance of monetary policy is measured by the use of monetary aggregates (the monetary base, *M1* and/or *M2*). This measure however is plagued by the *liquidity puzzle* (where positive innovations in these aggregates are associated with interest rate increases, contrary to what theory suggests) as documented by Leeper and Gordon(1992) as well as by Christiano *et al*(1992). Leeper and Gordon adopt a Wold causal interpretation of the data-set (on the rate of growth of the monetary base, $\Delta M0_t$; interest rates, R_t; consumer prices, P_t; and industrial production, Y_t; in the ordering given by $[\Delta M0_t, R_t, P_t, Y_t]'$) in which orthogonalized innovations in the rate of growth of the monetary base is considered a monetary policy shock. In a search for a solution to this puzzle Bernanke and Blinder(1992) as well as Sims(1992) identify monetary policy shocks directly with innovations in interest rates. However, quoting from Strongin(1995), "without any demonstrated empirical linkage between Federal Reserve actions and interest rate movements, it is unclear how innovations in interest rates can be reasonably attributed to monetary policy". It is this recognition that prompted Christiano et al(1992) - using various measures of money and of identification schemes (based on various Wold causal orderings of the data-set) to measure monetary policy shocks - to argue for the use of innovations in nonborrowed reserves as monetary policy shocks. But this implies, contrary to the empirical facts, that the Federal Reserve does not accommodate reserve demand shocks. The Strongin(1995) measure of innovations in the mix of borrowed and nonborrowed reserves as monetary policy shocks is therefore intended to address this conceptual shortcoming of Christiano et al(1992). Strongin's findings are not conclusive though. Bernanke and Mihov(1995) suggest a linear combination of innovations in total reserves, nonborrowed reserves and the federal funds rate as monetary policy shocks - implying that the Federal Reserve accommodates both borrowing and demand shocks in its monetary policy measures. Their identification scheme encompasses those of Bernanke and Blinder(1992), Christiano et al(1992) and Strongin(1995). We take this debate on the identification of monetary policy shocks a step further into the dynamics of exchange rates in response to monetary policy shocks. Our main concern here is with the issue of monetary policy interrelatedness and the extent to which an explicit modelling of the monetary policy arena of nations can help enhance our understanding of exchange rate fluctuations.

We follow the lead of Eichenbaum and Evans(1995) who use vector autoregressions (VAR) to analyze the effects of monetary policy shocks on bilateral exchange rate movements with the help of recent methods for identifying monetary policy shocks. We carry their analyses further in the light of the current state of the monetary policy innovations identification debate following Bernanke and Mihov(1995) who advocate a semi-structural VAR approach that does not only reflect the actual operating procedures of the federal reserve system but also nests all hitherto attempts - of Bernanke and Blinder(1992), Christiano and Eichenbaum (1992) and Strongin(1995) - at measuring monetary policy shocks. To address the pertinent issue of monetary policy interrelatedness and its effects on exchange rate fluctuations we introduce an exactly identified structural VAR model. Monetary policy interrelatedness is then explicitly modelled and tested, within the context of our VAR model, as over-identifying restrictions. Based on these two identification schemes we try to answer questions bordering on the extent to which earlier, rather unsatisfactory, empirical performance of the broad asset-market-based models of exchange rate determination can be attributed to the mode of identification of monetary policy shocks.

Using our identification schemes - one based on the actual conduct of monetary policy in the US and the other based on a structural VAR specification that explicitly incorporates international policy interrelatedness - we attempt to address the following questions/puzzles in the literature. Firstly, we investigate the extent to which our monetary policy identification schemes are able to solve the *forward discount bias puzzle*. Uncovered interest parity requires that positive domestic short-term interest rate innovations (relative to foreign short-term interest rate innovations) should be associated with an initial impact appreciation of the domestic currency; followed by a gradual depreciation of the domestic currency. But in the long-run a domestic currency appreciation vis-a-vis the US dollar is predicted - thus on impact, following the policy innovation, the exchange *overshoots* its long-run value. This is because, given uncovered interest parity, the positive interest rate differential has to be associated with an anticipation of a depreciation of the domestic currency in order for agents to continue holding foreign assets. Empirical findings not consistent with this requirement are said to yield a puzzle - the *forward discount bias puzzle*. This puzzle is encountered in Eichenbaum and Evans(1995) and Sims(1992) where (for some of the countries considered) positive domestic interest rate innovations are followed by large and persistent depreciations of the domestic currency. Secondly, could the *exchange rate puzzle* (which is the tendency of the domestic currency of the non-US G-7 countries to depreciate against the US dollar following contractionary domestic monetary policy shocks) that shows up in most empirical results in the literature be due to the specific schemes used to identify monetary policy shocks? This puzzle shows up in Sims(1992) as well as in Grilli and Roubini(1995) where - in a non-structural VAR approach - innovations in short term interest rates and monetary aggregates are used in the respective papers as monetary policy shocks. And finally, given our new identification schemes, we investigate the relative importance of the monetary policy shocks in explaining bilateral exchange rate fluctuations.

In an attempt to provide answers to the questions raised above we structure the paper as follows. The second section provides a brief review of the asset market based theories of exchange rate determination. This is meant to provide us with a reference theoretical framework that underlies subsequent empirical results. In the third section we briefly review the Bernanke-Mihov monetary innovations identification scheme and discuss our structural VAR model. The former approach - which we refer to in the rest of the paper as exemplifying the semi-recursive identification scheme - proposes innovations to an estimated linear combination of policy indicators as a measure of monetary policy shocks. The policy indicators used include total reserves normalised by the level of total reserves in the prior month, (TR_t), a measure of the mix of non-borrowed to total reserves (NBR_t) and the federal funds rate (FF_t). In section three we present and discuss our empirical results derived from estimating a seven-variable VAR in the logarithm of US industrial employment, the logarithm of the foreign country's industrial employment, the US total reserves, the US nonborrowed reserves, the US federal funds rate, a foreign monetary policy variable and the logarithm of the bilateral exchange rate (measured as the foreign country's currency per unit of the US dollar) respectively. The foreign countries considered are Canada, Germany, Japan and the UK. The estimated impulse response functions and variance decompositions of the logarithms of the nominal and real bilateral exchange rates are presented and discussed in the light of the general class of asset market models. The final section presents a summary of the approach adopted in the paper as well as of our empirical findings.

5.2 A Brief Review of Theory

In many standard textbooks in international finance - see for instance Obstfeld and Krugman(1994) or De Grauwe(1989) - the asset market based theoretical framework is introduced as a way to understanding exchange rate fluctuations. The monetary approach to exchange rate modelling encompasses both sticky-price models (of Dornbusch(1976) for instance) and flexible price models. One of the common elements of both of these variants of the monetary approach is that they are characterized by the implicit assumption of perfect substitutability between domestic and foreign money and bonds. In the vein of the latter model under this approach positive domestic monetary policy shocks lead to domestic price increases that depreciate the exchange rate given that purchasing power parity (PPP) holds in the short run. For the sticky-price models, where it is assumed that PPP does not hold in the short run, monetary policy operates through the interest rate channel. Thus positive domestic monetary policy shocks (implying an initial increase in the real money supply given that prices are sticky in the short run) reduce the domestic interest rate. Given uncovered interest parity and free capital mobility the policy shocks instigate an anticipation of an appreciation of the domestic currency in the long run. The fall in the domestic interest rate coupled with the expected appreciation of the domestic currency makes domestic assets unattractive. Domestic capital outflows - so long as the expected rate of appreciation of the domestic currency is greater than the interest rate gain (the interest rate differential) - and the domestic currency depreciates. A short run equilibrium is reached when equality is attained between the expected returns on foreign and domestic assets (i.e when uncovered interest parity holds). In the medium term however prices begin to rise as a result of the monetary expansion, the real money supply falls and domestic interest rate increases set in to gradually appreciate the domestic currency moving it to its long run purchasing power parity level. Thus according to this approach effects of positive monetary policy shocks are dominated in the short term by expectations and capital mobility in the asset market as well as by the exchange rate overshooting it's long run purchasing power parity level.

The second class of models under the broad class of asset market models - the portfolio balance model (PBM) - does not make the assumption of perfect substitutability between domestic and foreign financial assets. The PBM emphasizes wealth effects on asset demands and the role of the exchange rate and expectations about its future movements in the asset demand decisions. Some versions of the PBM introduce the current account balance in its role as allocating wealth among countries. A surplus/deficit in the current account represents a rise/fall in net domestic holdings of foreign assets which in turn affects the level of wealth and the real exchange rate. See for instance Branson(1983), and/or Taylor(1995) for an exposition of this approach. Following this approach the effects of monetary policy shocks are felt through agents' attempts to re-balance their portfolios by purchasing both foreign money and domestic bonds. Given that the supply of the domestic bonds is fixed (at least in the short run) the increase in its demand induces an increase in its price (i.e a reduction in the domestic interest rate) and depreciates the domestic currency.

Without explicitly discussing the dynamics of the exchange rate following a monetary policy shock we deem it sufficient for the purposes of this paper to present the expected shape of estimated impulse response functions of the exchange rate. Figures 5.1a and 5.1b below present the expected path of the exchange rate following US and foreign monetary tightening respectively.

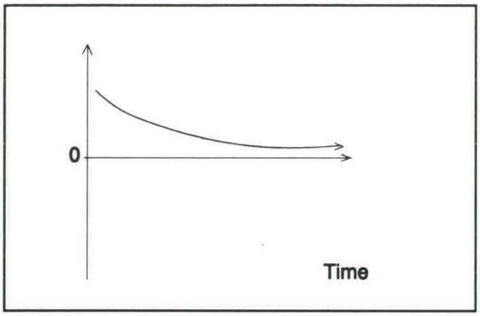


Figure 5.1a. The expected path of the exchange rate following US monetary tightening.

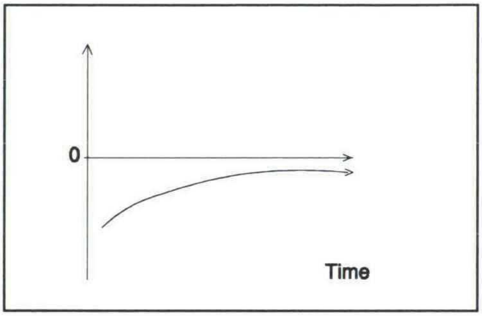


Figure 5.1b: The expected path of the exchange rate following foreign monetary tightening.

The graphs depicted in figures 5.1a and 5.1b above are based on the expected responses of the exchange rate on the basis of the overshooting phenomenon (under both theories discussed above) as well as from the predictions of uncovered interest parity under the sticky price model. One characteristic of these theories that comes out clearly in figure 5.1a is that, as a result of US monetary tightening, the US dollar appreciates on impact followed by a gradual depreciation. A corresponding expected path of the exchange rate following foreign monetary tightening is shown in figure 5.1b.

5.3 Identification Schemes

To state the problems in identifying monetary policy shocks, it is useful to briefly review the theory. Suppose the data generation mechanism for a covariance stationary X_t can (excluding deterministic components) be written as $G(L)X_t = v_t$, where $G(L)$ is a matrix lag polynomial of finite order, and v_t are economically meaningful, structural innovations, which we assume to be serially uncorrelated and to have a diagonal contemporaneous covariance matrix, $E[v_t v_t'] = \Xi$. If G_0 is invertible, the data generating mechanism can be written as a reduced form VAR, $C(L)X_t = u_t$, where $C_0 = I$ is the identity matrix of a suitable order and where $\Omega = E[u_t u_t'] = G_0^{-1}\Xi G_0^{-1'}$ is the covariance matrix of the one-step ahead prediction errors, u_t. Identification essentially involves reversing this process: using estimates of $C(L)$ and Ω to find $G(L)$ and Ξ. We shall use contemporaneous *a priori* exclusion restrictions on G_0 to identify the structural innovations. These restrictions are imposed such that the sum of $n(n+1)$ free parameters in G_0 and Ξ are reduced to no more than $n(n+1)/2$ distinct elements in Ω. To begin with n restrictions are obtained as normalizations that define the variances of the structural shocks by setting each of the n diagonal elements of G_0 to unity. For the remaining $n(n-1)/2$ restrictions various researchers have used different schemes. The fully recursive scheme implies the normalization assumption of a lower triangular G_0 matrix. The use of this normalization implies that variables pre-dating others are assumed to have contemporaneous effects on those coming after them but not vice versa. Thus the system is forced into taking on a contemporaneous recursive structure in a Wold causal fashion. We refer to this identification scheme as the fully recursive scheme in the rest of the paper.

Blanchard(1986) and Sims(1986) initiated a generalized method of allowing non-recursivity into the structure of G_0 and thereby allowing for more flexibility in modelling feedback effects among the set of variables and helps researchers impose more economic structure on the G_0 matrix thereby[1]. This way of identifying the G_0 matrix of contemporaneous effects has come to be known in the literature as the structural (or the identified) VAR approach as applied for instance by Bernanke and Mihov(1995)[2] in measuring monetary policy shocks in the US based on the actual operating procedures of the Federal Reserve System.

5.3.1 Monetary Policy Identification Schemes

A number of researchers have suggested various measures of monetary policy shocks as alternatives to the strategy of using orthogonalized innovations in such broad monetary aggregates as *M0, M1, M2* or *M3* as monetary policy shocks. These suggestions are put forward as attempts at solving the *liquidity puzzle*, the *exchange rate puzzle* and the *forward discount bias puzzle* encountered in earlier empirical work using innovations in these broad monetary aggregates as monetary policy shocks. Bernanke and Blinder(1992) argue for the use of innovations to the federal funds rate as monetary innovations. The analyses of Christiano et al(1994) as well as of Christiano and Eichenbaum(1992) argue for the use of innovations to nonborrowed reserves as monetary policy shocks whereas empirical evidence provided by

[1]. Apart from the exclusion restrictions generally favoured by researchers any set of restrictions can be imposed on the G_0 matrix to identify the structural innovations. Blanchard and Quah(1989) and W. D. Lastrapes(1992) use long-run restrictions whereas Faust and Leeper(1994) use a combination of both long run and short run restrictions to identify the underlying structural model.

[2]. More appropriately, the approach adopted by Bernanke and Mihov(1995) could also be classified as 'semi-structural' since they do identify their VAR by imposing restrictions derived from the actual operating procedures of the federal reserve system of the US (leaving the relationship among the non-policy variables unrestricted) and not just by the use of the fully recursive identification scheme under which G_0 is imply assumed to be lower triangular. However, to distinguish this approach from our structural VAR approach we denote the identification scheme exemplified by the Bernanke-Mihov(1995) approach as the semi-recursive identification scheme in the rest of the paper.

Strongin(1995) suggest innovations to a mix of borrowed and nonborrowed reserves. Extending the discussion further Bernanke and Mihov(1995) suggest a measure of monetary policy shocks that encompasses the earlier measures.

We present here only that part of Bernanke-Mihov that is relevant for our purposes - i.e the identification of monetary policy shocks. Bernanke and Mihov partition the matrices in the structural equation $G(L)X_t = v_t$ into the portion relating to a vector of non-policy macroeconomic variables, Y_t, and the part pertaining to the stance of monetary policy, P_t, such that $X_t = [Y_t, P_t]'$. Under the assumption that the contemporaneous portion of the policy authority's feedback rule contains information on the state of the economy (measured by the non-policy variables) we can derive the reduced-form relationships specified in equation (5.2) below for a given VAR(k) specification for X_t. The causal ordering adopted here, in X_t, implies that policy shocks do not have contemporaneous effects on the non-policy variables, and also that monetary policy authorities do have (and react to) contemporaneous information on the state of the economy. Given this specification a particular form of our general structural equation $G(L)X_t = v_t$ can be written as

$$\begin{aligned} Y_t &= \sum_{i=0}^{k} B_i Y_{t-i} + \sum_{i=1}^{k} H_i P_{t-i} + A^y v_t^y \\ P_t &= \sum_{i=0}^{k} D_i Y_{t-i} + \sum_{i=0}^{k} \Gamma_i P_{t-i} + A^p v_t^p \end{aligned} \qquad (5.1)$$

And the corresponding particular reduced form is derived as

$$Y_t = (I - B_0)^{-1}\left\{\sum_{i=1}^{k} B_i Y_{t-i} + \sum_{i=1}^{k} H_i P_{t-i} + A^y v_t^y\right\}$$

$$\begin{aligned} P_t = (I - \Gamma_0)^{-1}\Big\{\sum_{i=1}^{k}[D_i + D_0(I - B_0)^{-1}B_i]Y_{t-i} + \sum_{i=1}^{k}[\Gamma_i + D_0(I - B_0)^{-1}H_i]P_{t-i}] \\ + D_0(I - B_0)^{-1}A^y v^y + A^P v_t^P \Big\} \end{aligned} \tag{5.2}$$

where $v_t = [\ v_t^y,\ \ v_t^p\ ]'$ is a vector of unobserved structural innovations. More specifically the structural policy shocks vector v_t^p contains the elements v_t^d, v_t^b and v_t^s indicating total reserves demand shocks, borrowing shocks and non borrowed reserves supply shocks respectively. From the reduced-form above we infer that the component of observable innovations of P_t that are orthogonal to the non-policy vector can be indicated as

$$U_t^P = (I - \Gamma_0)^{-1} A^P v_t^P \tag{5.3}$$

The problem of identifying the stance of monetary policy involves the characterization and estimation of the matrix $(I - \Gamma_0)^{-1}A^p$. Bernanke and Mihov do this by using a simple conventional model of the bank reserve market as in Strongin(1995) but this time introducing borrowing shocks, v^b, as an additional element in nonborrowed reserves innovation (U_{NBR}). Viewing the federal funds rate as the opportunity cost of total reserves (the sum of required reserve and excess reserve holdings) held by banks they let innovations in the demand for reserves (U_{TR}) react negatively to innovations in the federal funds rate (U_{FF}). The rest of the model describing the behaviour of the federal reserve is as specified in the set of equations (5.4) below

$$
\begin{aligned}
U_{TR} &= -\alpha U_{FF} + \nu^{d} \\
U_{BR} &= \beta U_{FF} + \nu^{b} \\
U_{NBR} &= \phi^{d}\nu^{d} + \phi^{b}\nu^{b} + \nu^{s}
\end{aligned}
\tag{5.4}
$$

where α, β, ϕ^d and ϕ^b are parameters to be estimated whereas ν^d, ν^b and ν^s are as defined above. From the set of equations in *(5.4)* above we can solve for the observable policy innovations ($U^p = [U_{FF}, U_{TR}, U_{NBR}]'$) as functions of borrowing, demand and policy shocks - ν^b, ν^d and ν^s - respectively as in equation (5.5) below using the identity, $TR = NBR + BR$.

$$
\begin{bmatrix} U_{TR} \\ U_{FF} \\ U_{NBR} \end{bmatrix} = \begin{bmatrix} -\dfrac{\alpha}{\alpha+\beta}(1-\phi^{d})+1 & \dfrac{\alpha}{\alpha+\beta} & \dfrac{\alpha}{\alpha+\beta}(1-\phi^{b}) \\ \phi^{d} & 1 & \phi^{b} \\ \dfrac{1}{\alpha+\beta}(1-\phi^{d}) & -\dfrac{1}{\alpha+\beta} & -\dfrac{1}{\alpha+\beta}(1+\phi^{b}) \end{bmatrix} \begin{bmatrix} \nu^{d} \\ \nu^{s} \\ \nu^{b} \end{bmatrix}
\tag{5.5}
$$

This equation contains seven unknowns (if we include the variances of the structural shocks) to be estimated from six covariances between the observable policy innovations, U^p. Thus we need *at least* one identifying assumption to be able to pin down the monetary policy shock, ν^s which is a combination of the elements of the policy innovations matrix U^p. In Strongin(1995) it is argued that the Fed accommodates reserve demand shocks through the provision of nonborrowed reserves or through the discount window so that shocks to total reserves reflect only changes in the demand for reserves at least in the short run. This implies, in the context of model in equation (5.5) above, that α and ϕ^b be set to zero. The Bernanke-Mihov exactly identifying assumption which yields the monetary policy shock as a linear combination of innovations in total reserves, nonborrowed reserves and the federal funds rate after setting only α to zero is specified in equation (5.6) below.

$$v^s = -(\phi^d + \phi^b)U_{TR} + (1 + \phi^b)U_{NBR} + \beta\phi^b U_{FF} \qquad (5.6)$$

Notice that by setting $\phi^b = 0$ into equation (5.6) above we obtain the Strongin(1995) scheme where monetary policy shocks are identified as a mix of innovations in total and nonborrowed reserves. Bernanke and Mihov also show that the identification schemes of Bernanke and Blinder(1992) as well as that of Christiano *et al.*(1992) can be derived as special cases of their general model for which α is non-zero[3]. We adopt Bernanke-Mihov's exactly identified model - as an example of the semi-recursive identification scheme - in the analysis of exchange rate fluctuations in response to monetary policy shocks in subsequent sections of this paper.

5.3.2 International Policy Interdependence

There are certain relevant issues bordering on the contemporaneous feedback structure of our application of the Bernanke-Mihov identification scheme - issues of international monetary policy interdependence and of the feedback from exchange rates to monetary policy formulation and vice versa. To address these issues we introduce a structural VAR that explicitly incorporates policy interdependence and the reaction of monetary policy authorities to exchange rate movements as verifiable hypotheses. We then compare the performance of

[3]. Notice that for a non-zero α equation (5.6) can be replaced by

$$v^s = -(\phi^d + \phi^b)U_{TR} + (1 + \phi^b)U_{NBR} - (\alpha\phi^d - \beta\phi^b)U_{FF} \qquad (5.6')$$

and the required restrictions in the case of the Bernanke-blinder specification of attributing innovations in interest rates to monetary policy are for $\phi^d = 1$ and $\phi^b = -1$. The Christiano et al case requires setting $\phi^d = \phi^b = 0$ implying that innovations in nonborrowed reserves represent monetary policy shocks.

the semi-recursive identification scheme, as exemplified by the Bernanke-Mihov scheme, with that of our structural VAR approach. For both identification schemes we use the seven variable VAR with the variables as defined earlier on. We follow the lead of most researchers in this area of international finance and use VARs that imply asymmetric treatment of the non-US countries vis-a-vis the US. This could be justified by the fact that monetary institutions and the conduct of monetary policy do differ between countries. Hence there is not much reason to expect innovations in a linear combination of total reserves, nonborrowed reserves and the federal funds rate (or some other relevant interest rate) to adequately measure monetary policy shocks in the non-US countries just because this metric measures monetary policy adequately well in the US. In Canada the relevant monetary policy indicators that adequately reflect the operational behaviour of the monetary authority are excess settlement cash and the overnight rate. In fact Armour *et al*(1996) provide empirical evidence to the effect that the innovations in the overnight rate adequately reflects the stance of monetary policy in Canada. To empirically investigate the assertion above we apply the Bernanke-Mihov approach to German data. We run a five variable VAR using monthly data on the logarithm of industrial production, the consumer price index, total reserves, nonborrowed reserves and the day-to-day interest rate (tagesgeld) over the period 1974:1 - 1994:9. The empirical results, based on Bernanke-Mihov's exactly identified model (that is the model with $\alpha=0$ in equation (5.5); implying that innovations in total reserves are independent of the funds rate) yields a *liquidity puzzle* where positive monetary policy shocks are associated with significant increases in the interest rates. Hence, based on the empirical evidence cited above and also to escape this *liquidity puzzle* that we do not directly address in this paper, we use innovations to short term interest rates in the non-US countries as monetary policy shocks.

Considering the high degree of interdependence among economies through trade and capital movements; as well as the fact that exchange rate fluctuations have much influence on domestic economic activity (especially in the highly open economies that we consider here) we deem it informative to introduce international monetary policy interdependence explicitly into the model[4]. The dependence of domestic monetary policy on foreign macroeconomic variables and policy could be attributable to a common reaction to global shocks. Conscious efforts by policymakers to stabilize their respective currencies on the global financial market are also fertile grounds for this type of interdependence among nations. We favour the second channel in this paper. The issue of international transmission of domestic monetary policy is not new in the literature. In an open economy, with a flexible exchange rate system, unanticipated money supply shocks are associated with movements of short-term interest rates and bilateral exchange rates because of changes in interest parity relationships after the policy announcement. The divergence from covered interest parity so created can only be attained by a movement in the forward rate or of the foreign short-term interest rate. The interdependences that we consider here are those among policy shocks and non-policy variables. We specifically categorize these interdependences as follows:

i. the impact of the exchange rate on US monetary policy and vice versa,
ii. the impact of the exchange rate on foreign monetary policy and vice versa, and finally
iii. the impact of US monetary variables on foreign monetary policy and vice versa.

The necessary restrictions (= $n(n-1)/2$) for exactly identifying our structural VAR are introduced as exclusion restrictions on the contemporaneous structural parameters in G_0. Following Bernanke(1986) and our discussion of interdependence above these restrictions yield a particular form of our G_0 matrix (we refer to it as $\mathbf{\Phi}$ in the rest of the paper) as summarized in the set of relationships in equation (5.7) below.

[4]. The idea of considering policy interdependence in empirical policy analysis in the context of open economies is not completely new. It has been demonstrated in the monetary policy interdependence literature (see Sephton(1989) for instance) that due to structural linkages among economies monetary policy authorities can attain trend-stationarity in the price level and exchange rates by optimally conditioning their policy rules on foreign monetary policy innovations.

$$\begin{bmatrix} u_{emp}^{us} \\ u_{emp}^{f} \\ u_{tr} \\ u_{nbr} \\ u_{ff} \\ u_{m}^{f} \\ u_{xr} \end{bmatrix} = \begin{bmatrix} 1 & 0 & 0 & 0 & 0 & 0 & 0 \\ g_{21} & 1 & 0 & 0 & 0 & 0 & 0 \\ g_{31} & 0 & 1 & 0 & 0 & 0 & 0 \\ g_{41} & 0 & g_{43} & 1 & g_{45} & g_{46} & g_{47} \\ g_{51} & 0 & g_{53} & g_{54} & 1 & g_{56} & g_{57} \\ 0 & g_{62} & 0 & g_{64} & 0 & 1 & g_{67} \\ g_{71} & g_{72} & g_{73} & g_{74} & g_{75} & g_{76} & 1 \end{bmatrix} \begin{bmatrix} v_{us}^{y} \\ v_{f}^{y} \\ v^{d} \\ v^{s} \\ v^{b} \\ v_{f}^{s} \\ v_{x} \end{bmatrix} \tag{5.7}$$

Our structural model above identifies monetary policy in the foreign country and in the US with innovations in short term interest rates and the ratio of nonborrowed reserves to total reserves respectively. Following Bernanke and Blinder(1992) we assume that monetary policy variables respond contemporaneously to innovations in domestic macroeconomic variables - this is to say that monetary authorities have contemporaneous information on the state of the economy. The sixth equation in our system of equation (5.7) refers to the foreign monetary authority's contemporaneous feedback rule for setting r^F. Foreign monetary authorities react to contemporaneous information on the exchange rate and the stance monetary policy in the US. The contemporaneous policy interaction parameter measuring the feedback from US monetary policy to the foreign country's policy stance is g_{64}. Similarly US monetary policy is assumed to react contemporaneously to the stance of foreign monetary policy and the exchange rate with the interaction parameters given by g_{46} and g_{47} respectively. This specification - that the US reacts to foreign monetary policy and vice versa - is worth some empirical scrutiny. We test the particular over-identifying restriction implied by relaxing any one of our assumptions pertaining to policy interrelatedness in turns and condition subsequent estimations and analyses on the results of the tests.

The interdependence that appear in our structural model above is the key difference between our identification scheme and those of researchers using the fully recursive schemes and semi-recursive schemes (as exemplified by Strongin (1995) and Bernanke and Mihov(1996)). We think that once we introduce international aspects of monetary policy formulation into the debate of measuring monetary policy shocks the issues that are of the most interest are those of interdependence between policy and non-policy variables as outlined above. These interdependences are not addressed in the semi-recursive scheme of Bernanke and Mihov. Also, given the high degree of international interdependence among nations through the effects of exchange rate fluctuations on trade and capital movements, any monetary policy identification scheme aimed at explaining exchange rate fluctuations should necessarily incorporate this interdependence. The fully recursive scheme as used in Eichenbaum and Evans(1995) implies allowing all non-policy variables to influence all policy variables, but fails to address the relevant issues of interdependence as listed above. Of particular significance in this paper is the impact of our structural identification scheme to the dissolution of the exchange rate and forward discount bias puzzles encountered in the literature.

The rest of our identifying restrictions are as follows. We assume that monetary policy shocks have no contemporaneous effects on non-policy variables[5]. The last line of our contemporaneous feedback matrix regards the exchange rate as a forwardlooking asset price and hence models exchange rate shocks as a combination of innovations in all the variables in our model.

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[5]. This assumption is perhaps the only common ground between our structural model, the fully recursive and semi-recursive identification schemes that we consider in this paper. The variable orderings adopted by Bernanke and Blinder(1992), Bernanke-Mihov(1995), Christiano *et al*(1992) and Eichenbaum and Evans(1995) - to mention but a few - imply this very assumption.

5.4 Empirical Results

The empirical evidence presented below are based on seven variable (S)VARs estimated using monthly data for Canada, Germany, Japan and the UK over the period 1974:01 to 1994:05. A lag length of twelfth months is used in each case. The choice of the lag length is based on evidence regarding the non-existence of serial correlation in the error terms as measured by the Ljung-Box Q statistics as in Doan(1992). Insofar as the error terms in our estimated (S)VARs behave normally we think that testing for co-integration and basing subsequent estimations on the cointegration vector, if any, would not yield any significantly different results in terms of the impulse responses. Neither would the use of first differences of the data yield any significant differences in our estimates. The data are presented and described in appendix A. For each country the (S)VARs include, as endogenous variables, the logarithm of employment in the United States and in the country being considered (i.e. $log.\ emp_t^{US}$ and $log.\ emp_t^{F}$ respectively) measured in weekly hours of work in manufacturing, US total reserves TR_t, US nonborrowed reserves NBR_t, US federal funds rate FF_t, the foreign country's short-term interest rate, r_t^F and the logarithm of the nominal bilateral exchange rate $S_t^{(/\$)}$ measured as the foreign country's (domestic) currency price of one US dollar. When we include the logarithm of the real bilateral exchange rate instead of the logarithm of the nominal the former is measured as the difference between $S_t^{(/\$)}$ and the logarithm of the ratio of the US consumer price index to the foreign country's. Defined this way an increase in $S_t^{(/\$)}$ implies an appreciation of the dollar (or equivalently, a depreciation of the foreign currency). We include eleven monthly seasonal dummies as well in the (S)VARs.

We present three sets of empirical evidence from our estimations. Our first set of empirical evidence is for a fully recursive scheme using the ordering as in

$X_t = \left[\; log.emp_t^{US} \; log.emp_t^{F} \; TR_t \; NBR_t \; FF_t \; r_t^{F} \; s_t^{(/\$)} \; \right]'$. The second set of results is based on the semi-recursive identification scheme as exemplified by the Bernanke-Mihov monetary policy identification scheme. The third set of empirical results is based on our structural identification scheme where we incorporate policy interrelatedness by way of exclusion restrictions on the contemporaneous reactions matrix, Φ. Preliminary tests for over-identifying restrictions based on the assumptions either of US policy independence of foreign policy shocks and vice-versa for each of the countries considered yield very low levels of statistical significance. Hence, under our structural VAR identification scheme, we estimate seven variable structural VAR based on the exactly identified model presented in equation (*5.7*) explicitly incorporating the assumption of international monetary policy interdependence.

To regenerate the exchange rate puzzle and the forward discount bias puzzle we use the fully recursive identification scheme since it is the most common scheme applied in this area of research. The Wold causal ordering used in this experiment is the same as that in X_t above. This particular ordering implies, among others, that included in the US monetary policy reaction function are domestic and foreign economic conditions; but not the foreign monetary policy variable. Secondly it implies the assumption that foreign monetary authorities' incorporate information on US domestic economic conditions and the stance of US monetary policy in formulating their policies. And finally, the exchange rate is assumed to be affected contemporaneously by innovations in all the variables in the system.

Our empirical results as presented in figure 5.2a and 5.2b are based on the fully recursive identification scheme. The estimated impulse responses show that a positive monetary policy shock in the US leads to a fall in the federal funds rate and depreciates the dollar (except in the case of the UK) over the entire horizon of our impulse response experiments. However, from the estimated impulse responses reported in the first column of figure 5.5b for the exchange rate, a contractionary foreign monetary policy (measured as a positive innovation to the foreign short term interest rate) generally leads to an impact appreciation of the dollar vis-a-vis the foreign currency over the first five months following the policy shock. Thus the triangular identification scheme results in the exchange rate puzzle encountered in the literature. This evidence suggests the need for a re-examination of the data using more flexible identification schemes that accord some weight to international policy interdependence. Grilli and Roubini(1995) suggest however that the exchange rate puzzle could be blamed on identification schemes that imply the assumption of the US as a 'leader' - and all other countries are 'followers' - in the international policy arena. They argue further that short term interest rate innovations in non-US G-7 countries are endogenous policy reaction to inflationary shocks that in themselves cause domestic currency depreciations. Our empirical results cast some doubts on the plausibility of their explanations of the source of the puzzle however. As we show below using the semi-recursive identification structure that implies treating the US as an international policy 'leader' and the rest of the world as 'followers' the exchange rate puzzle largely disappears - even in the presence of the 'leader-follower' characterization. This finding suggests that the use of the fully recursive scheme for identifying monetary policy shocks is one likely source of the puzzles. Secondly, to the extent that policymakers are aware of the negative implications of domestic inflationary shocks on the domestic currency, it seems rather unlikely that they would not incorporate this knowledge (of the reaction of exchange rates to domestic inflation) into the formulation of policy.

5.4.1 Measuring Policy Innovations and Policy Interdependence

The estimates of some of the parameters of our exactly-identified seven-variable structural VAR are presented in table 5.1 below[6]. We utilize these estimates in investigating the extent to which our estimated monetary policy shocks display characteristics that qualify them as such. Generally, on *a priori* theoretical grounds, we expect an expansionary monetary policy shock in the US to lead to a fall in the federal funds rate and - following the predictions of asset-market-based models of exchange rate determination (the overshooting sticky-price model of Dornbusch(1976) for instance)-to, *ceteris paribus*, depreciate the dollar on impact just as a contractionary foreign monetary policy shock would. To investigate the extent to which our identified innovations can be regarded as plausibly representing what one would expect from the normal operation of the federal funds market we present and discuss the estimated impact effects of innovations in nonborrowed reserves on the federal funds rate - i.e the estimates of g_{54}. Estimates of the respective contemporaneous effects of innovations in US nonborrowed reserves and in the foreign short term interest rate on the exchange rate - i.e g_{74} and g_{76} respectively. To fix some ideas about international policy interrelatedness we also present and discuss the estimates of g_{64} in the light of theoretical predictions. Two entries are presented under each estimated parameter in table 5.1 below. The first entry shows the estimates from our seven variable structural VAR in X_t with the bilateral nominal exchange rate as the last element whereas the second entry is when the bilateral real exchange rate is used[7].

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[6]. We present only a subset of the estimates here in order to concentrate on the parameters of direct interest to us in discussing issues of policy identification as well as of international policy interrelatedness. All the $n(n-1)/2=21$ parameter estimates are reported in table B1 in appendix B.

[7]. For the Bernanke-Mihov identification scheme as used in this paper the parameters of interest include β, ϕ^b and ϕ^d. The estimates of these parameters are 0.0323, 0.7125 and -0.0299 respectively for Canada; 0.0414, 0.6362, and 0.0860 respectively for Germany; 0.0339, 0.7115 and 0.0036 for Japan; and finally 0.0362, 0.7229 and -0.0025 for the UK. The estimates for β indicate, as predicted by theory, that increases in the federal funds rate (i.e. the rate at which borrowed reserves are traded in at the federal funds market) increases banks' demand for borrowed reserves - see equation (*5.4*) above. These estimates are obtained for VARs including nominal bilateral exchange rates - the estimates are not significantly different from these when we use real bilateral exchange rates.

Table 5.1: ***Estimates of contemporaneous policy/exchange rate interaction parameters***

	Parameter Estimates							
Country	g_{54}		g_{64}		g_{74}		g_{76}	
Canada	-1.095*	-0.149*	-0.226	-1.697*	-0.014	-0.099*	-0.007*	-0.006*
	(-)	(-)	(0.477)	(-)	(0.274)	(-)	(-)	(-)
Germany	-0.508*	-13.782*	- 7.202*	-0.130	-0.118*	-0.147*	-0.013*	-0.013*
	(-)	(-)	(1.225)	(0.857)	(-)	(-)	(-)	(-)
Japan	-0.206	-0.184*	1.965	-0.065	-0.142*	-0.269*	-0.003*	-0.008*
	(0.488)	(-)	(2.691)	(0.791)	(-)	(-)	(-)	(-)
UK	-0.167*	-0.142*	-0.206	-0.214	-0.110	-0.113*	-0.002	-0.001
	(-)	(-)	(0.675)	(0.501)	(0.378)	(-)	(0.009)	(0.005)

Notes: *The estimates presented in this table are based on our exactly-identified structural identification scheme. The standard errors of the estimates in parentheses. A * indicates significance at the 95% confidence level. The estimates are obtained using the Simplex Algorithm and the results used as initial values in the Bernanke procedure as in the RATS 4.0 manual. Figures indicated by " - " are standard errors less than 1×10^{-3}.*

From the table presented above we deduce the following. First, we advance sufficient evidence in support of the use of our identified structural innovations to nonborrowed reserves, v^{US}_{NBR}, as US monetary policy shocks. Secondly, we investigate the relationship between these shocks and the policy innovations of the respective foreign countries. Finally, we investigate the extent to which the domestic and foreign policy shocks so identified have the desired (impact) effects on the exchange rate. To address the first issue it is worth noting that for all countries and for both cases considered expansive innovations in nonborrowed reserves significantly reduce the federal funds rate on impact. In addition to this evidence the estimate of g_{53} (the impact effect of total reserves innovations on the federal funds rate) as presented in table 5B of appendix 5B is generally significantly negative indicating that demand shocks do increase the federal funds rate on impact. Then comes the issue of international policy interdependence. Using estimated innovations in the foreign short term interest rate as foreign policy shocks yields the plausible result that foreign monetary policy shocks respond, on immediate impact, negatively to a positive US monetary policy innovation - except for Japanese data with the nominal exchange rate ordered last. This empirical finding together with the results of our formal tests of policy interdependence, as discussed so far in this section, suggest the existence of some form of policy interrelatedness between the US on the one hand and each of the countries considered on the other. A second way of verifying the plausibility of our estimated monetary policy shocks is to examine whether their impact effects on bilateral exchange rates are in accordance with a priori theoretical predictions. From the estimates of g_{74} and g_{76} given in the table above the US dollar depreciates on impact following expansionary US monetary policy shocks and positive foreign interest rate shocks respectively. These results indicate that, given our structural VAR identification scheme, it is reasonably justifiable to consider the estimated linear combination of innovations in total reserves, the federal funds rate, the exchange rate and foreign monetary policy as monetary policy shocks in the US. For the non-US countries considered here monetary policy shocks in this case can be reasonably represented by a linear combination of short term interest rate innovations, US nonborrowed reserves innovations and exchange rate innovations.

5.4.2 Dynamic Responses

The foreign exchange market and the international policy arena - including those of strategic monetary policy aimed at exchange rate stabilization as considered in this paper - are best understood within a dynamic setting. In this section therefore we present an analysis of our estimates of the dynamic effects of monetary policy shocks on real and nominal bilateral exchange rates using estimated impulse responses and forecast error variance decompositions based on the identification schemes discussed above. We also examine the extent to which these schemes help explain the puzzles encountered in the literature.

Estimated Impulse Responses and Variance Decompositions

Based on the semi-recursive identification scheme, figure 5.3a depicts the estimated responses for the respective countries following a one standard deviation US monetary policy shock. Figure 5.3b presents equivalent estimates of the responses to a one standard deviation foreign monetary policy shock. Figures 5.4a and 5.4b are the exact parallels of figures 5.3a and 5.3b except that in this case the estimated impulse responses are based on our structural identification scheme. Rows 1 to 4 of figures 5.2 through 5.4 show the estimated responses obtained using Canadian, German, Japanese and British data respectively. The columns 1 through 5 depict the estimated response of the US nonborrowed reserves, US federal funds rate, the foreign short term interest rate, the nominal exchange rate, and the real exchange rate respectively to a one-standard-deviation monetary policy shock as indicated in the title. The nominal exchange rate is measured as the domestic currency price of a unit of the US dollar. To conserve space we do not present the estimated responses for the logarithm of employment in the figures mentioned above. Figures 5.5a and 5.5b assist us in the discussion of the estimated exchange rate responses as well as in assessing the relative performance of the various identification schemes considered in helping solve or generate the puzzles. The rows of figure 5.5a refer to the estimated responses of the nominal exchange rate to a one standard deviation foreign monetary policy shock obtained using Canadian, Japanese, German and British data respectively. The columns correspond to the various identification schemes used in estimating the responses. Figure 5.5b presents a corresponding information except that in this case the estimated responses are with respect to a one standard deviation foreign contractionary monetary policy shock. Two-standard-error bands of the estimated impulse responses over 48 months time span - shown as dashed lines in the figures referred to above - are calculated by generating 500 random draws from the estimated asymptotic distribution of the VAR coefficients and the covariance matrix of the innovations[8]. The x-axis of each of the estimated responses shows the time horizon over which the impulse response experiments are performed.

Considering the estimated effects of expansive monetary policy disturbances in the US, as presented in figures 5.3a and 5.4a, a number of important results emerge. Firstly, the negative response of the federal funds rate to US monetary policy shocks shows up under both identification schemes and for all countries considered. Secondly, it turns out that generally, as the federal funds rate falls following expansive monetary policy in the US, the short term interest rate of the respective foreign countries considered here also falls significantly. We interpret this as an indication of policy interrelatedness. Since foreign monetary authorities' monetary reaction function includes the US federal funds rate, as US interest rates fall foreign monetary authorities intervene to mitigate the envisaged subsequent fall in the dollar (alternatively an appreciation of their domestic currency) and the train effects of worsening trade balance. Finally, an expansive monetary policy in the US, depreciates the dollar - in real and nominal terms - on impact. The depreciation is however not as persistent under the structural identification scheme as it is for the semi-recursive identification scheme. Given our structural identification scheme, except for Canada and Japan, we do not observe persistent depreciation of the dollar following US expansive monetary policy shocks. In fact, from the estimated impulse responses as in figure 5.5a, the minimum level of depreciation of the dollar, following US monetary policy shocks, occurs a little after a month in the case of Germany and the UK. However, whereas the minimum depreciation is attained eventually (after roughly 25

[8]. In the case of exactly identified models, unlike in the case of over-identified VAR models, there is a one-to-one mapping of draws from the set of possible reduced form models into a set of possible structural interpretations. Hence our estimated error bands - calculated using the Monte Carlo procedure packaged with the RATS software - cannot be improved upon using the corrective adjustments suggested in Sims and Zha(1994). Killian(1995) establishes, using Monte Carlo simulations, that a bias-corrected bootstrap interval outperforms the Monte Carlo interval as well as the standard bootstrapping interval and the asymptotic interval in accounting for the bias observed in estimated small-sample impulse responses. His arguments however are valid only in the context of exactly identified *bi-variate* structural VAR models however; and hence are of no consequence for our results.

months for the nominal exchange rate and 9 months for the real exchange rate) for Canadian data, Japanese data exhibits persistence in the depreciation of the dollar. Thus, in terms of consistency with the hypothesis of uncovered interest parity and overshooting models of exchange variability following monetary policy shocks, the structural identification scheme (incorporating policy interdependence) performs relatively better than the semi-recursive scheme. There is, under the latter scheme, a general tendency towards persistent depreciation of the dollar following expansionary US monetary policy shocks. This is generally the case under the fully recursive identification scheme as well.

Next, consider figures 5.3b, 5.4b and 5.5b which report corresponding results following contractionary foreign monetary policy shocks. Observe the *exchange rate puzzle* distinctly exhibited under the fully recursive identification scheme shown in the first column of figure 5.5b.. Generally, for the semi-recursive and structural identification schemes and for all the countries considered foreign contractionary monetary policy shocks lead to an impact depreciation of the dollar (or equivalently, an appreciation of the domestic currencies of these countries). Except for the case of Japan the pattern looks very similar in all the countries under the semi-recursive scheme. On the other hand our structural identification scheme again yields results consistent with the predictions of uncovered interest parity under which the increase in the domestic interest rate following domestic monetary tightening must be associated with a domestic currency appreciation upon impact followed by a gradual depreciation. The impact dollar depreciations, following foreign short term interest rate shocks, are followed at least by gradual appreciations of the dollar. More specifically, following contractionary foreign monetary policy shocks the minimum depreciation of the dollar occurs after the first month in Japan. Again, in consonance with the hypothesis of uncovered interest parity. The process is however gradual for the rest of the countries considered - the minimum depreciation rate occurs after 10, 9 and 15 months for Canada, Germany, and the UK respectively. Thus even though the expected depreciations of the domestic currency do not occur right after the impact appreciations, our results are far from the persistent appreciations for periods of up to two years observed by earlier researchers.

We turn next to assessing the relative importance of the identified monetary policy shocks - under each identification scheme - in explaining exchange rate fluctuations. Table 5.2 below summarizes the estimated proportions (in percentages) of forecast error variances of the exchange rate over the indicated horizons (indicated in months). In the exception of a few instances where exchange rate innovations account for less than 50% of the forecast error variances the largest proportion of the forecast error variance is generally explained by exchange rate innovations themselves.

Further, the results indicate that the estimated forecast error variance decompositions vary quite substantially not only across countries and schemes but also between the two monetary policy shocks. Aspects of the general picture as depicted by the table above are summarized in the couple of points below:

i) US monetary policy shocks generally explain a relatively larger proportion of the forecast error variance of the bilateral exchange rates than the foreign monetary policy shocks do. In particular, under our structural identification scheme, US monetary policy shocks explain roughly 31.06% of the forecast error variance for Canadian data whereas foreign monetary policy shocks explain only about 8.92% over a 12 month horizon. The corresponding figures for German and Japanese data are {51.17% and 0.91%} and {41.75% and 11.01%} respectively. Given the semi-recursive identification scheme the figures drop to barely {3.28% and 0.50%} for German data and {11.49% and 1.92%} for Japanese data. This evidence is however not supported by the data-set for the UK under both identification schemes where the relative explanatory power is switched in favour of foreign monetary policy innovations.

Table 5.2: Estimated Forecast Error Variance Decompositions for Nominal Exchange Rates

Horizon	Shocks	Semi-Recursive Identification Scheme				Structural Identification Scheme			
		Canada	Germany	Japan	UK	Canada	Germany	Japan	UK
1	v^{US}_{NBR}	3.04	4.01	4.75	-	29.99	50.93	20.85	0.31
	v^{F}_{M}	7.10	2.49	0.14	0.21	9.11	0.78	1.60	64.88
	v_{XR}	85.84	84.65	91.51	96.23	57.14	47.47	39.47	24.13
2	v^{US}_{NBR}	4.02	5.41	8.46	-	30.94	51.04	24.69	0.48
	v^{F}_{M}	4.89	1.65	0.11	1.87	8.64	0.82	1.26	60.41
	v_{XR}	85.31	79.84	88.49	90.74	56.64	47.33	39.85	28.31
4	v^{US}_{NBR}	3.47	5.91	8.42	0.44	30.81	51.03	25.01	1.54
	v^{F}_{M}	3.93	0.89	1.05	5.45	8.99	0.83	0.91	52.23
	v_{XR}	86.09	74.55	88.03	80.06	56.39	47.34	38.12	36.09
6	v^{US}_{NBR}	5.78	5.25	9.10	0.34	31.98	51.02	30.11	3.82
	v^{F}_{M}	3.47	0.65	1.49	7.40	8.48	0.82	2.28	44.28
	v_{XR}	81.43	72.79	84.95	71.17	55.49	47.37	33.89	40.50
12	v^{US}_{NBR}	5.61	3.28	11.49	0.39	31.06	51.17	41.75	7.25
	v^{F}_{M}	1.78	0.50	1.92	17.08	8.92	0.91	11.01	35.80
	v_{XR}	77.41	70.38	76.10	51.41	55.78	47.17	20.15	42.97

Notes: The numbers presented in this table are in percentages. A ' - ' is used to represent numbers less than 1×10^{-3}.

ii) Generally the sum of the proportion of the forecast error variance explained by both US and foreign monetary policy shocks is perceptibly higher for our structural identification scheme than for the semi-recursive scheme. Again, this is an indication that the former identification scheme incorporating international monetary policy interdependence performs better that the scheme that regards the US as a policy 'leader' and all other countries as 'followers'.

Based on the evidence presented so far the more flexible structural identification scheme that explicitly incorporates policy interdependence among countries can be said to encompass the semi-recursive identification scheme to a very large extent.

5.5 Summary and Conclusion

Earlier empirical research on the effects of exchange rate fluctuations in response to monetary policy shocks encountered a number of puzzles including the *exchange rate puzzle* and the forward discount bias puzzle. This paper introduces current findings in the debate in monetary economics - on the measurement of monetary policy shocks - into an open economy context. The main objective here is to investigate the extent to which the documented puzzles could be attributable to the particular monetary policy identification schemes used. The main finding of this paper is that the puzzles could be attributable to the use of the fully recursive identification scheme in VARs applied in the open economy. The low degree of flexibility of the fully recursive identification scheme rarely delivers models that reflect the complexities of the arena of policy-making. We regard international policy interdependence as one of these complexities that the recursive identification scheme is unable to capture. In particular, the scheme, as is well-known, restricts the reaction functions of monetary authorities in such a way that such necessary factors as exchange rate and foreign monetary policy variables cannot enter the policy reaction function of the US while at the same time allowing these same variables to react contemporaneously to US monetary policy shocks. This non-flexibility is addressed in this paper by introducing a structural identification scheme that allows us to incorporate contemporaneous exchange rate movements and international policy interrelatedness into the monetary authorities' reaction functions.

As the empirical results reveal the structural identification schemes used in this paper are able to add to our understanding of the effects of monetary policy shocks on exchange rate fluctuations and yield results consistent with *a priori* theoretical predictions. One of these identifications schemes - called the semi-recursive identification scheme in the paper - is based on the actual operating procedures of the US federal reserve system. The second scheme follows the structural VARs tradition by isolating a set of independent shocks by way of theoretically meaningful restrictions In particular our empirical findings suggest that indeed monetary policy identification schemes, especially those incorporating international policy linkages, could help solve the puzzles.

5.6 References

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Appendix 5A: *Data Sources and Description*

The data used in the estimations are monthly data obtained from OECD MEI Database for the period 1974:1 to 1994:12 (unless otherwise indicated). The various variables include the following for the respective countries.

1. Employment (measured in weekly hours of work in manufacturing unless otherwise stated):
   a) Canada ..: Canhours
   b) Germany...: Deumhour
   c) United Kingdom...: Gbrhours
   d) Japan...: Jpnhours
   e) United States...: Usahours

2. Monetary policy indicators/variables:
   a) Canada ..: Canprime (Prime Interbank Rate - Commercial Banks)
   b) Germany...: Deucall (the German Call Money Rate)
   c) United Kingdom...: Gbrcall (Call Money Rate)
   d) Japan...: Jpngbond (Central Gov't Bond Yields)
   e) United States...: TR (Total Reserves - CITIBASE Data)
   NBR (Nonborrowed Reserves - CITIBASE Data)
   FF (Federal Funds Rate - CITIBASE Data)

3. Exchange rates: (All exchange rates used are indicated as domestic currency per unit of the US dollar; monthly averages).
   a) Canada...: Canusxau
   b) Germany...: Deuusxav
   c) United Kingdom...: Gbrusxav
   d) Japan...: Jpnusxav

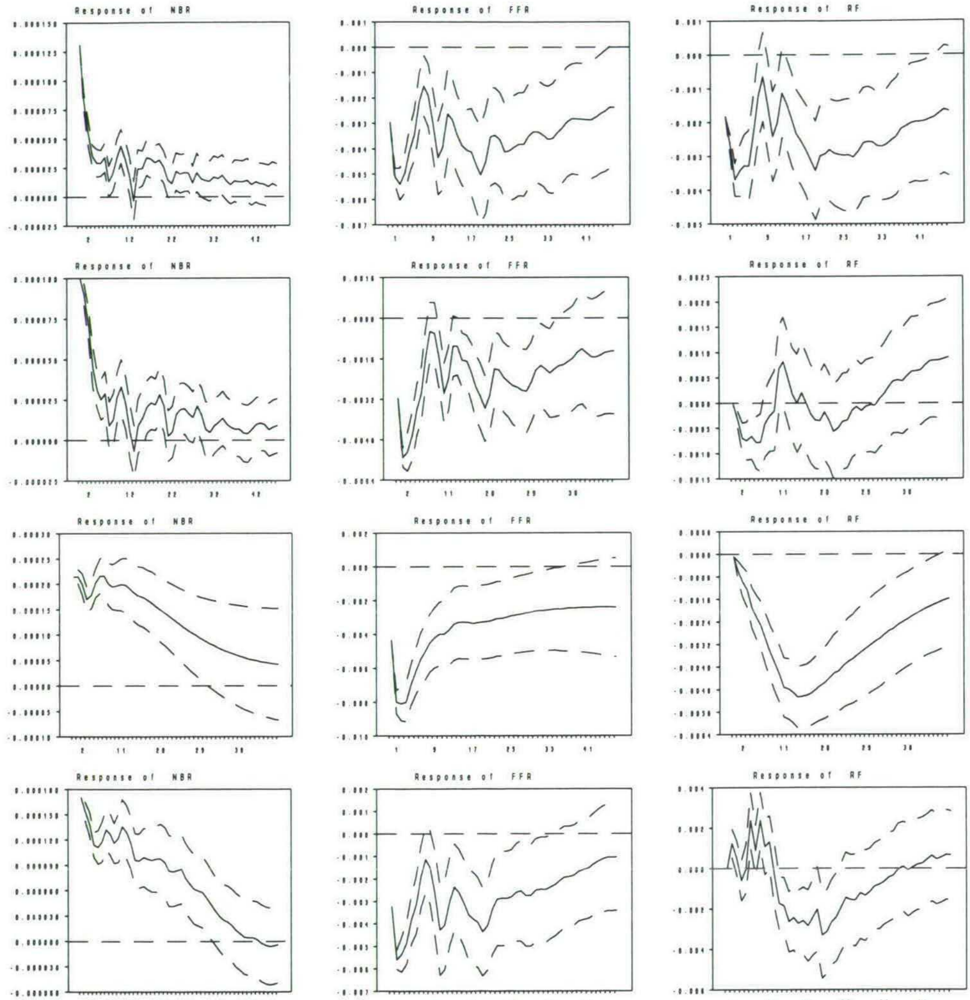
Appendix 5B: *Parameter Estimates of the Structural VAR model*

Table 5B *Parameter Estimates of the Structural VAR model*

Parameter Estimates	Country							
	Canada		Germany		Japan		UK	
g_{21}	-0.173 (0.048)	-0.177 (0.048)	0.021 (0.372)	0.142 (0.372)	-1.574 (0.222)	-1.466 (0.217)	0.035 (0.059)	0.025 (0.059)
g_{31}	0.185 (0.088)	0.189 (0.087)	0.061 (0.083)	0.078 (0.083)	-0.079 (0.075)	-0.086 (0.074)	0.021 (0.076)	0.009 (0.077)
g_{41}	-0.083 (0.068)	-0.072 (-)	-0.139 (0.099)	-0.135 (0.070)	-0.087 (0.064)	-0.076 (-)	-0.065 (0.063)	-0.063 (0.071)
g_{43}	-0.852 (0.063)	-0.811 (-)	-0.833 (0.079)	-0.820 (0.072)	-0.832 (0.103)	-0.828 (-)	-0.849 (0.097)	-0.872 (0.061)
g_{45}	0.015 (-)	0.014 (-)	0.013 (-)	0.016 (-)	0.012 (0.005)	0.012 (-)	0.013 (-)	0.013 (-)
g_{46}	-0.001 (0.001)	0.000 (-)	0.012 (0.004)	0.003 (0.004)	-0.001 (0.003)	0.000 (-)	-0.001 (0.001)	-0.001 (0.001)
g_{47}	0.092 (0.348)	0.072 (-)	0.033 (-)	0.028 (-)	0.034 (-)	0.043 (-)	0.016 (0.274)	0.035 (-)
g_{51}	-0.006 (0.946)	-0.015 (-)	-16.106 (3.996)	-0.001 (-)	-2.640 (0.689)	-0.275 (0.724)	-0.003 (0.789)	-0.061 (0.016)
g_{53}	-0.015 (-)	0.012 (-)	-0.095 (-)	-0.015 (-)	-6.905 (0.206)	-0.034 (-)	-0.003 (-)	-0.003 (-)
g_{54}	-1.095 (-)	-0.149 (-)	-0.508 (-)	-13.782 (-)	-0.206 (0.488)	-0.184 (-)	-0.167 (-)	-0.142 (-)
g_{56}	-0.664 (-)	-0.669 (-)	-0.389 (-)	-0.011 (0.077)	0.102 (0.264)	-0.043 (-)	-0.066 (0.062)	-0.036 (0.053)
g_{57}	-1.106 (-)	-0.848 (-)	-1.063 (0.598)	-1.254 (-)	-1.273 (0.403)	-0.985 (-)	-1.303 (-)	-1.036 (-)
g_{62}	0.002 (0.582)	-0.007 (0.870)	-0.091 (-)	0.516 (0.515)	-2.499 (0.640)	-0.001 (0.569)	0.000 (0.827)	-0.005 (0.672)
g_{64}	-0.226 (0.477)	-1.697 (-)	-7.202 (1.225)	-0.130 (0.857)	1.965 (0.691)	-0.065 (0.791)	-0.206 (0.675)	-0.214 (0.501)
g_{67}	0.174 (0.439)	-4.272 (-)	-0.311 (-)	0.019 (-)	0.011 (0.789)	-0.003 (-)	1.769 (0.454)	0.018 (0.181)
g_{71}	0.032 (0.065)	0.003 (-)	0.036 (-)	0.052 (-)	0.221 (0.144)	0.227 (-)	0.011 (0.143)	0.018 (0.262)
g_{72}	0.009 (0.067)	-0.022 (-)	-0.038 (0.030)	-0.039 (-)	-0.001 (0.043)	0.004 (-)	0.189 (0.058)	0.017 (0.025)
g_{73}	-0.022 (0.230)	0.041 (-)	0.045 (-)	0.083 (-)	0.006 (-)	-0.001 (-)	-0.148 (1.115)	-0.086 (-)
g_{74}	-0.014 (0.274)	-0.099 (-)	-0.118 (-)	-0.147 (-)	-0.142 (-)	-0.269 (-)	-0.110 (0.378)	-0.113 (-)
g_{75}	0.001 (0.006)	0.000 (-)	-0.011 (-)	-0.013 (-)	-0.008) (-)	-0.009 (-)	-0.001 (0.014)	-0.002 (0.012)
g_{76}	-0.007 (-)	-0.006 (-)	-0.013 (-)	-0.013 (-)	-0.003 (-)	-0.008 (-)	-0.002 (0.009)	-0.001 (0.005)

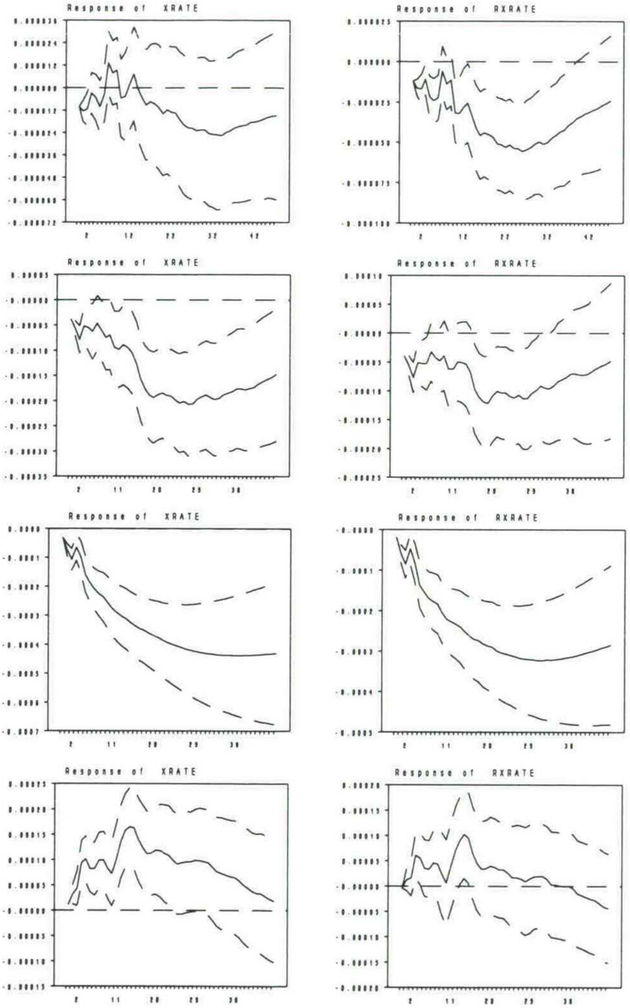
Notes: *The estimates in parentheses are standard errors. Two elements are presented for each country. The first entry shows the estimates from a VAR including the bilateral nominal exchange rate as the last element whereas the second entry is when the bilateral real exchange rate is used. The estimates are obtained using the Simplex Algorithm and the results used as initial values in the Bernanke procedure as in the RATS 4.0 manual. Figures indicated by "-" are less than 1×10^{-3}. We use 12 lags in all the estimations.*

Figure 5.2a. Estimated impulse responses to US monetary policy shocks under the fully recursive identification scheme



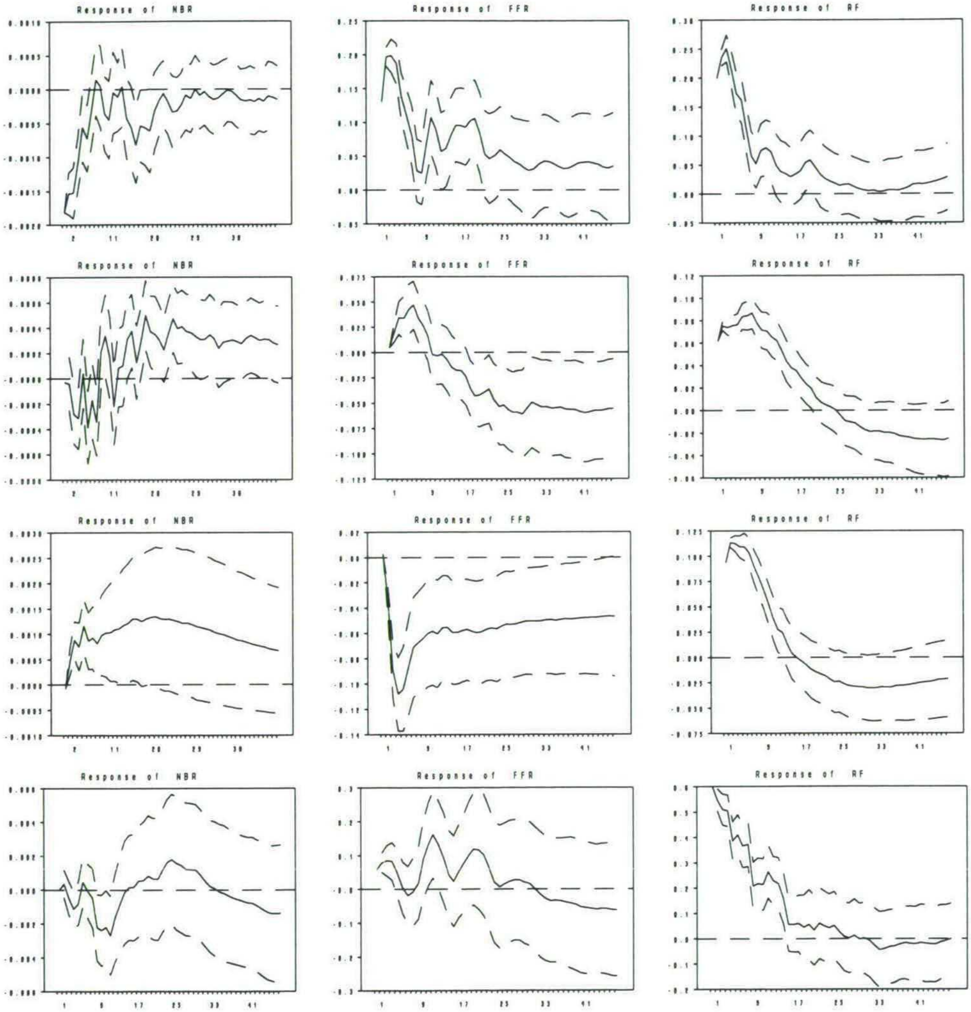
Notes: Rows 1 to 4 show the estimated responses obtained using Canadian, German, Japanese and British data respectively. Columns 1 through 3 depict the estimated response of the US nonborrowed reserves, US federal funds rate and the foreign short term interest rate respectively to a one-standard-deviation US monetary policy shock.

Figure 5.2a. (contd.): ***columns 4 & 5***



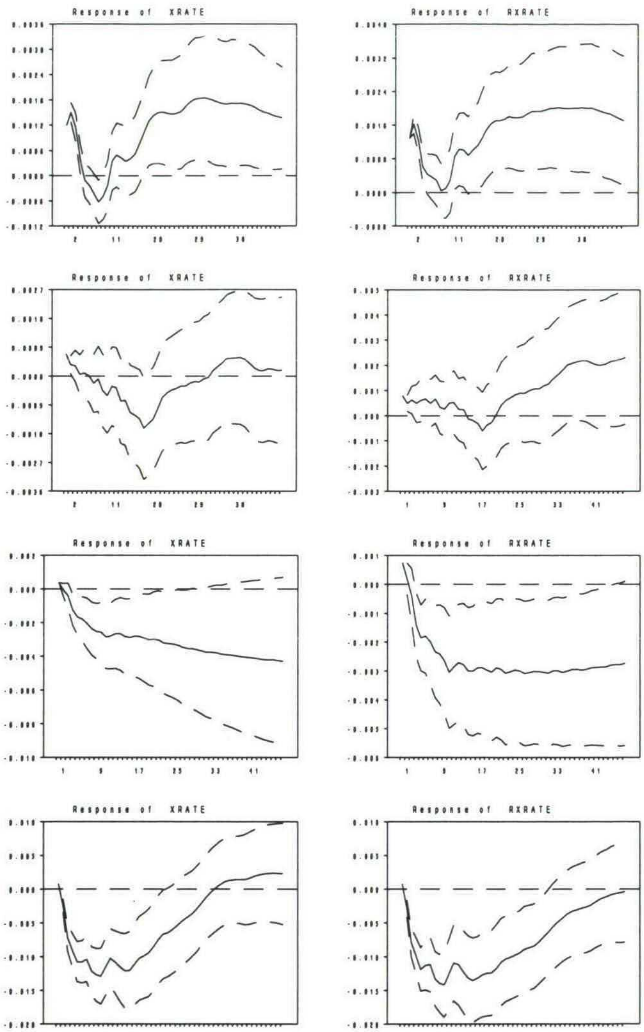
Notes: Rows 1 to 4 show the estimated responses obtained using Canadian, German, Japanese and British data respectively. The columns 4 and 5 depict the estimated response of the nominal exchange rate, and the real exchange rate respectively to a one-standard-deviation US monetary policy shock. The nominal exchange rate is measured as the domestic currency price of a unit of the US dollar.

Figure 5.2b. Estimated impulse response to foreign monetary policy shocks under the fully recursive identification scheme



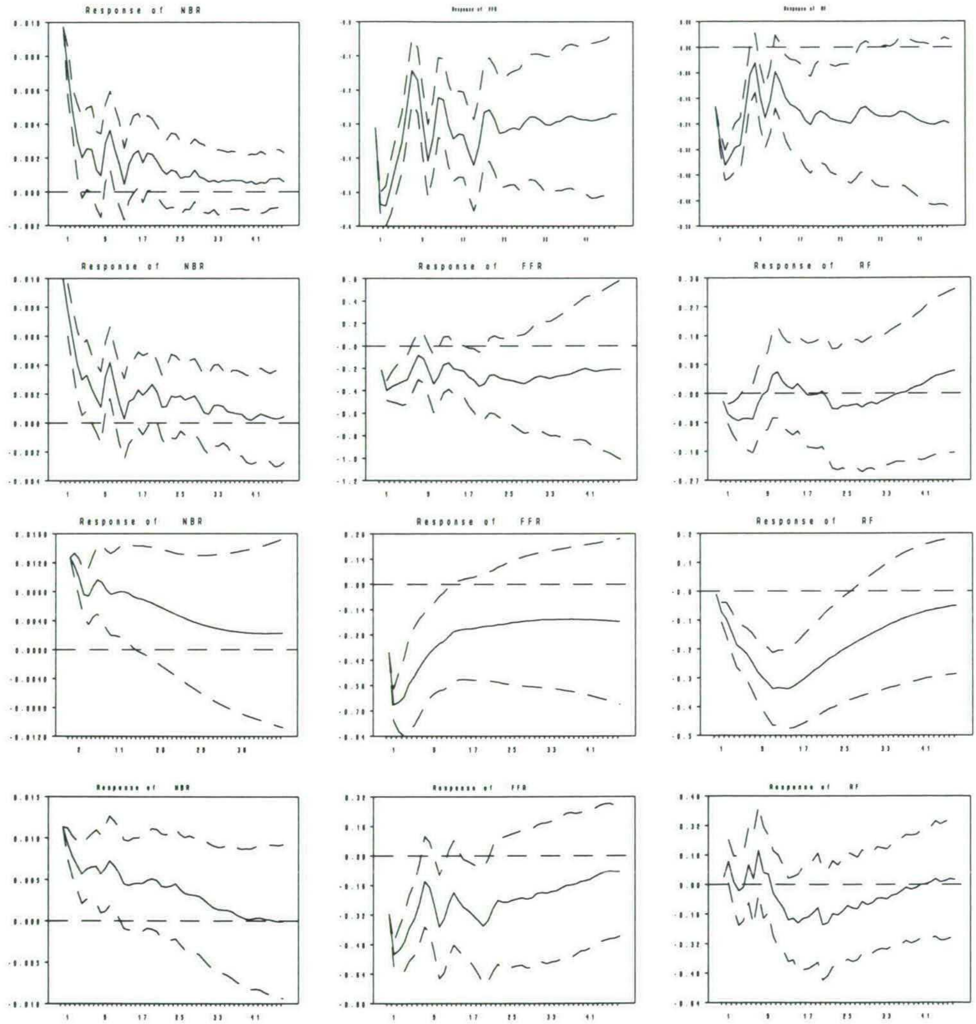
Notes: Rows 1 to 4 show the estimated responses obtained using Canadian, German, Japanese and British data respectively. Columns 1 through 3 depict the estimated response of the US nonborrowed reserves, US federal funds rate, the foreign short term interest rate, the nominal exchange rate, and the real exchange rate respectively to a one-standard-deviation foreign monetary policy shock.

Figure 5.2b. (contd.): *columns 4 & 5*



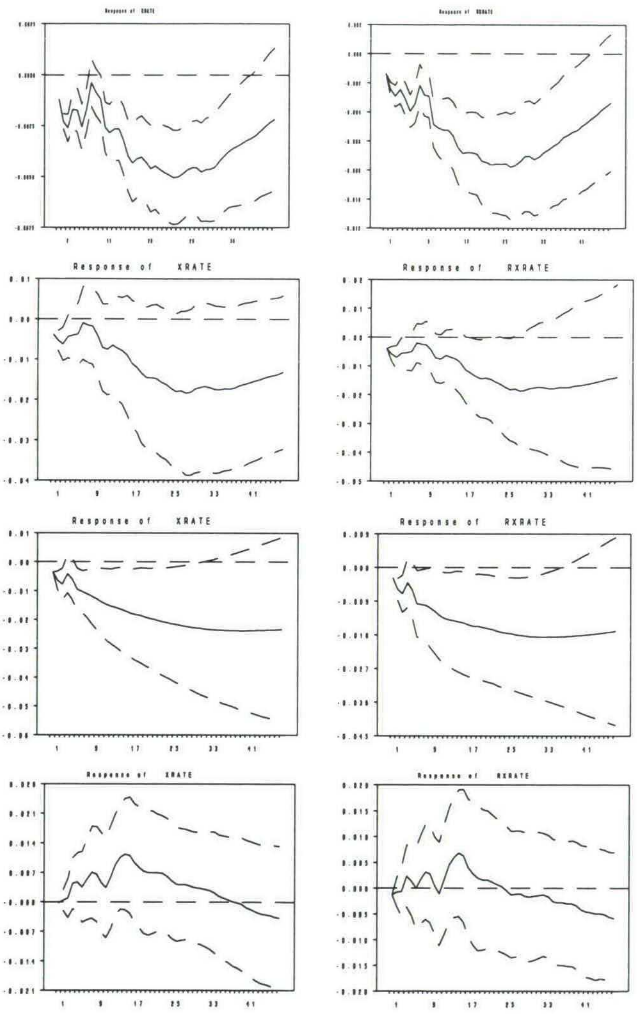
Notes: Rows 1 to 4 show the estimated responses obtained using Canadian, German, Japanese and British data respectively. The columns 4 and 5 depict the estimated response of the nominal exchange rate, and the real exchange rate respectively to a one-standard-deviation foreign monetary policy shock. The nominal exchange rate is measured as the domestic currency price of a unit of the US dollar.

Figure 5.3a. Estimated impulse responses to US Monetary Policy Shocks; given the semi-recursive identification scheme



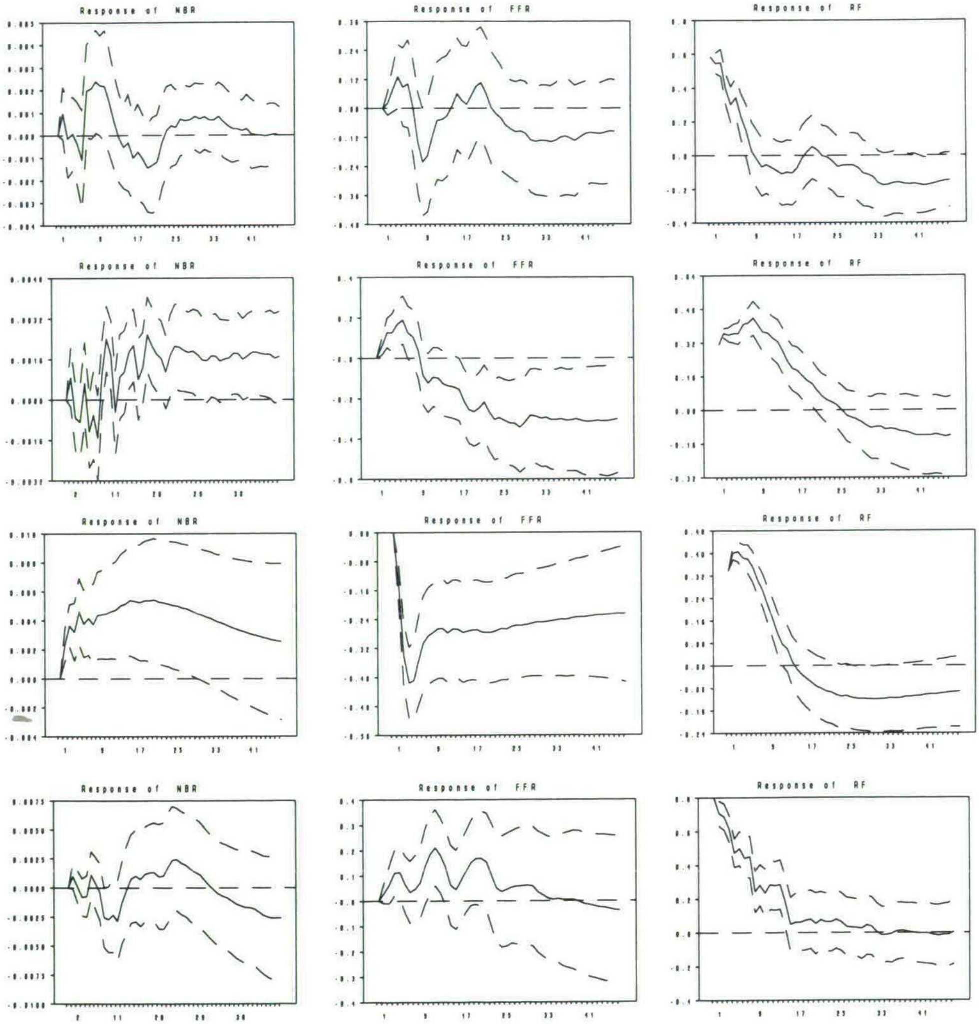
Notes: Rows 1 to 4 show the estimated responses obtained using Canadian, German, Japanese and British data respectively. Columns 1 through 3 depict the estimated response of the US nonborrowed reserves, US federal funds rate and the foreign short term interest rate respectively to a one-standard-deviation US monetary policy shock.

Figure 5.3a. (Contd.): ***columns 4 & 5***



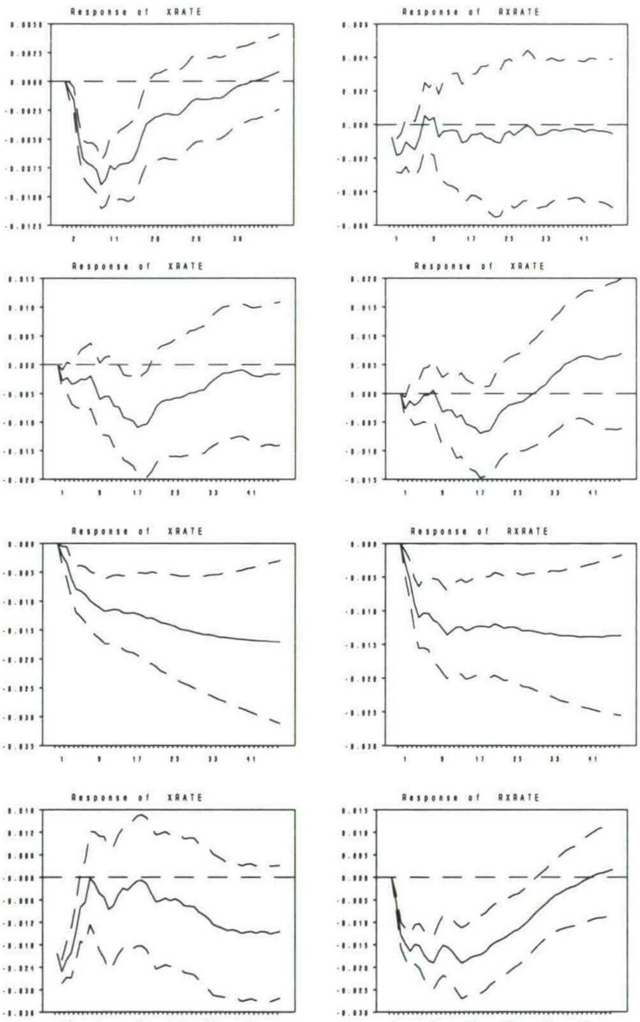
Notes: Rows 1 to 4 show the estimated responses obtained using Canadian, German, Japanese and British data respectively. The columns 4 and 5 depict the estimated response of the the nominal exchange rate, and the real exchange rate respectively to a one-standard-deviation US monetary policy shock. The nominal exchange rate is measured as the domestic currency price of a unit of the US dollar.

Figure 5.3b. Estimated impulse responses to foreign monetary policy shocks; given the semi-recursive identification scheme



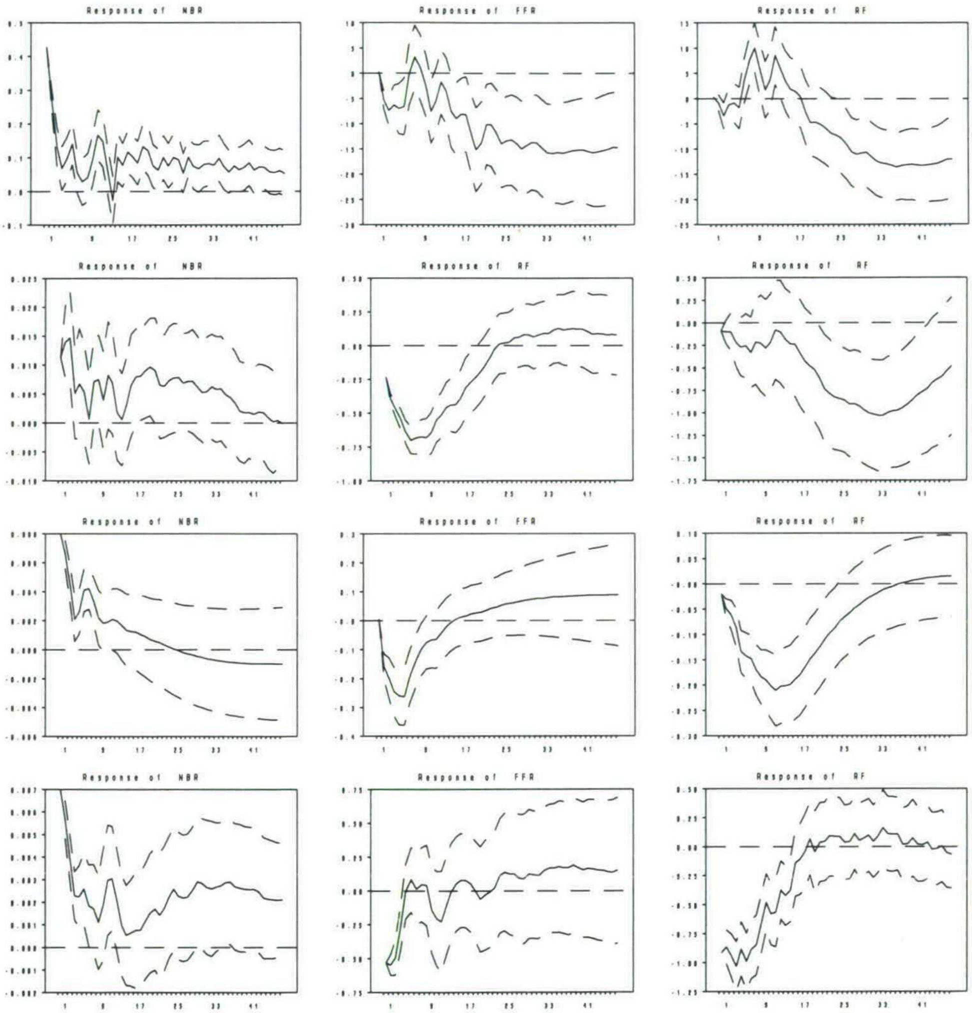
Notes: Rows 1 to 4 show the estimated responses obtained using Canadian, German, Japanese and British data respectively. Columns 1 through 3 depict the estimated response of the US nonborrowed reserves, US federal funds rate, the foreign short term interest rate respectively to a one-standard-deviation foreign monetary policy shock.

Figure 5.3b. (Contd.): *columns 4 & 5*



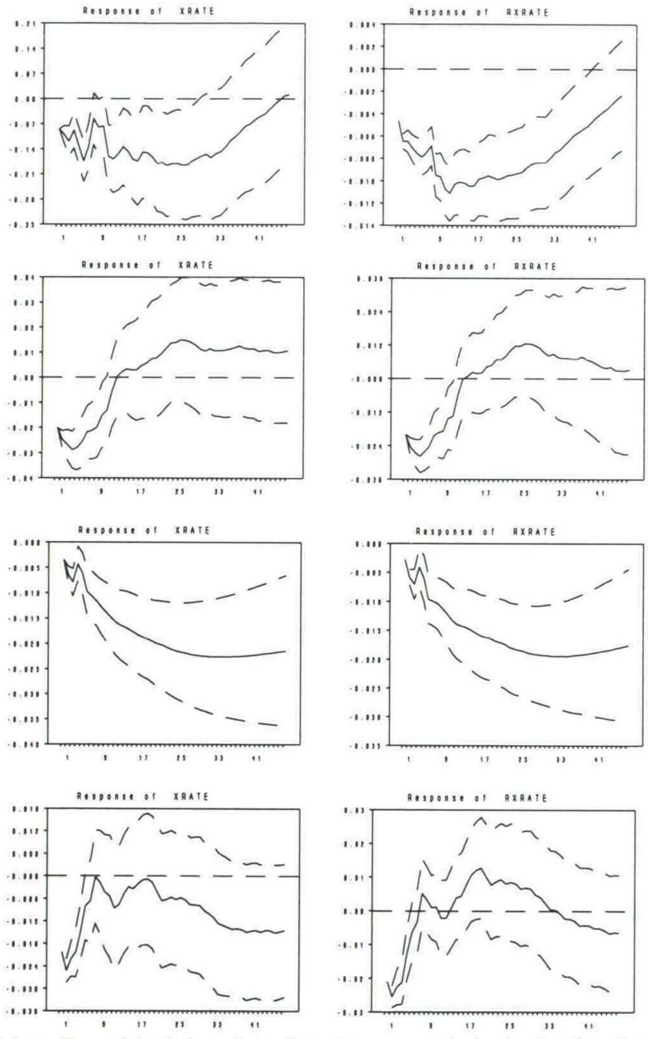
Notes: Rows 1 to 4 show the estimated responses obtained using Canadian, German, Japanese and British data respectively. The columns 4 and 5 depict the estimated response of the nominal exchange rate and the real exchange rate respectively to a one-standard-deviation foreign monetary policy shock. The nominal exchange rate is measured as the domestic currency price of a unit of the US dollar.

Figure 5.4a. Estimated impulse responses to US Monetary policy shock; given the structural identification scheme



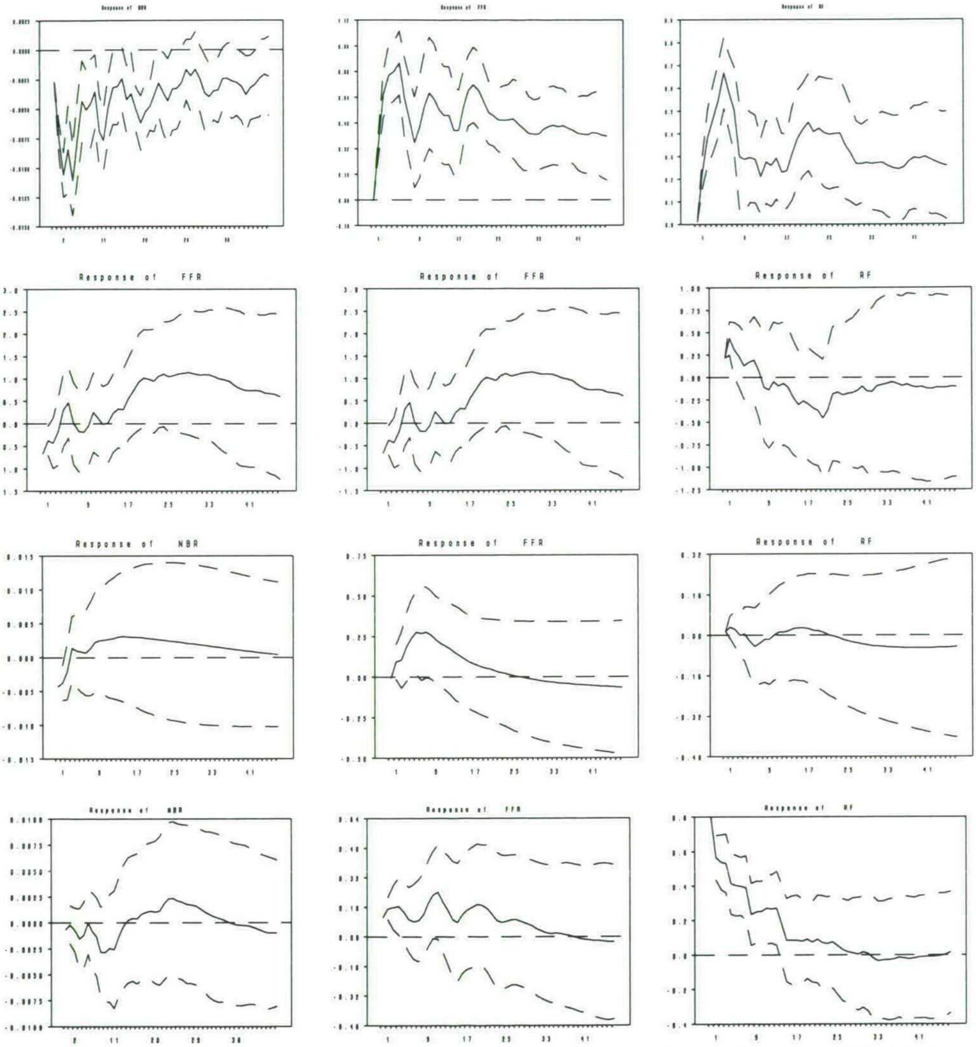
Notes: Rows 1 to 4 show the estimated responses obtained using Canadian, German, Japanese and British data respectively. Columns 1 through 3 depict the estimated response of the US nonborrowed reserves, US federal funds rate and the foreign short term interest rate respectively to a one-standard-deviation US monetary policy shock.

Figure 5.4a. (Contd.): ***columns 4 & 5***



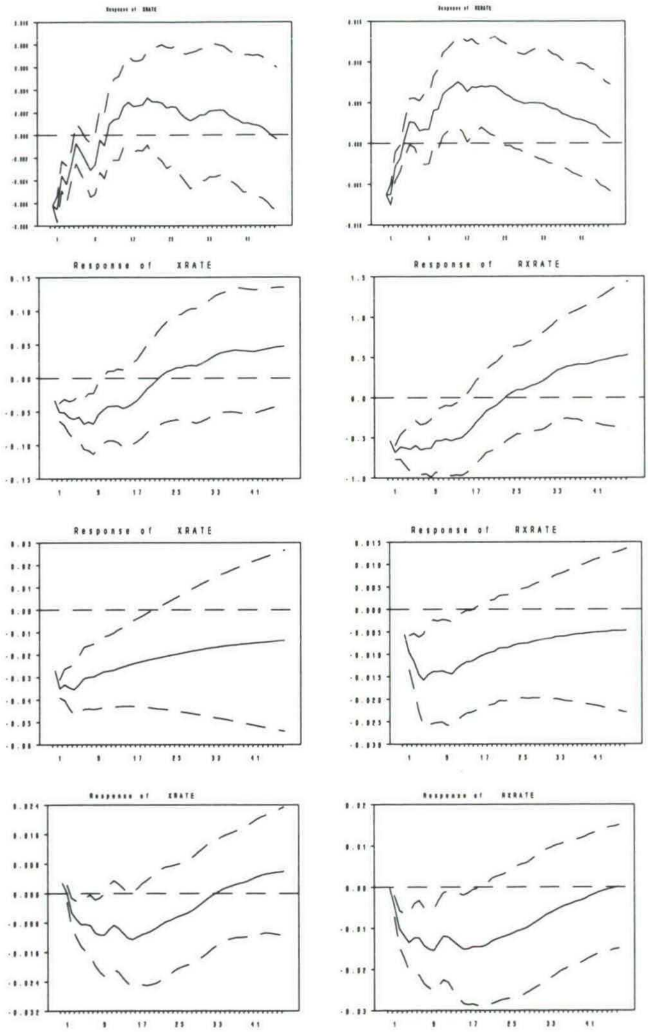
Notes: Rows 1 to 4 show the estimated responses obtained using Canadian, German, Japanese and British data respectively. The columns 4 and 5 depict the estimated response of the nominal exchange rate, and the real exchange rate respectively to a one-standard-deviation US monetary policy shock. The nominal exchange rate is measured as the domestic currency price of a unit of the US dollar.

Figure 5.4b. Estimated impulse responses to foreign policy shocks; given the structural identification scheme



Notes: Rows 1 to 4 show the estimated responses obtained using Canadian, German, Japanese and British data respectively. Columns 1 through 3 depict the estimated response of the US nonborrowed reserves, US federal funds rate and the foreign short term interest rate respectively to a one-standard-deviation foreign monetary policy shock.

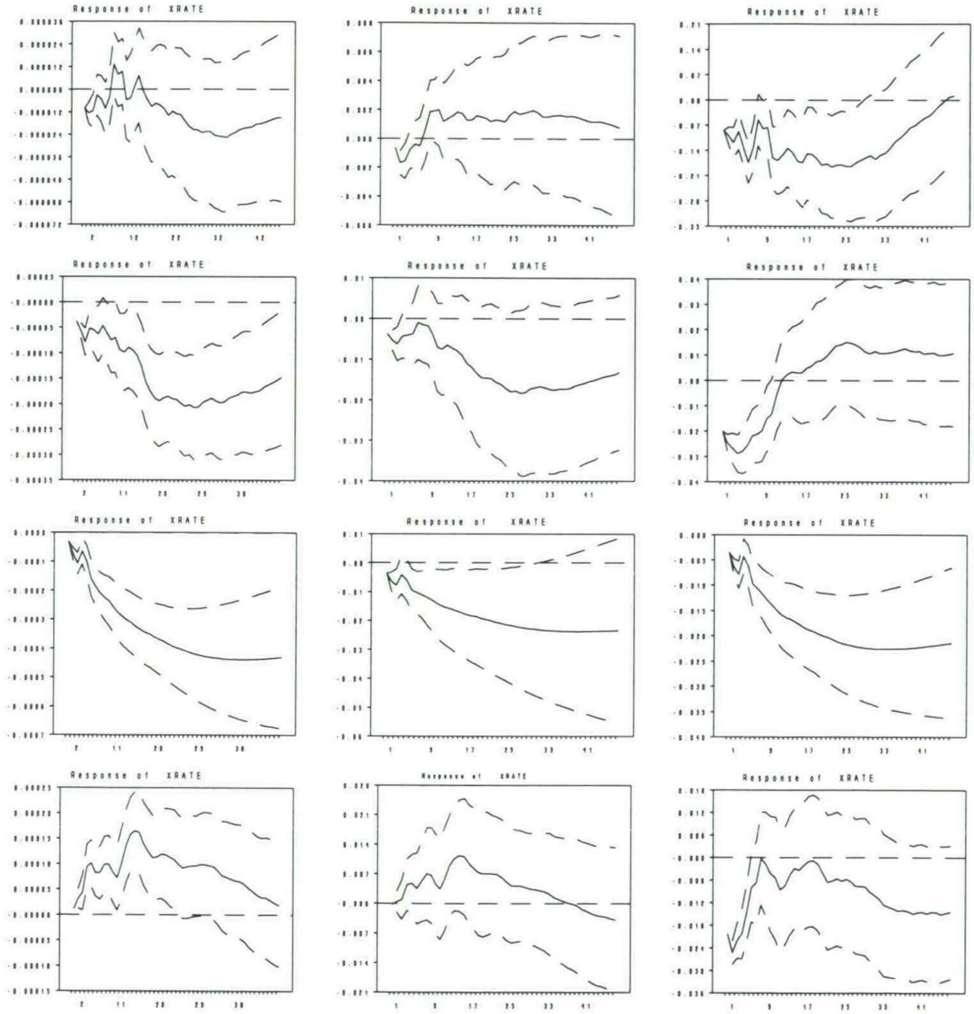
Figure 5.4b. (Contd.): ***columns 4 & 5***



Notes: Rows 1 to 4 show the estimated responses obtained using Canadian, German, Japanese and British data respectively. The columns 4 and 5 depict the estimated response of the nominal exchange rate, and the real exchange rate respectively to a one-standard-deviation foreign monetary policy shock. The nominal exchange rate is measured as the domestic currency price of a unit of the US dollar.

Figure 5.5a. Monetary policy identification schemes and estimated exchange rate responses to US monetary policy shock

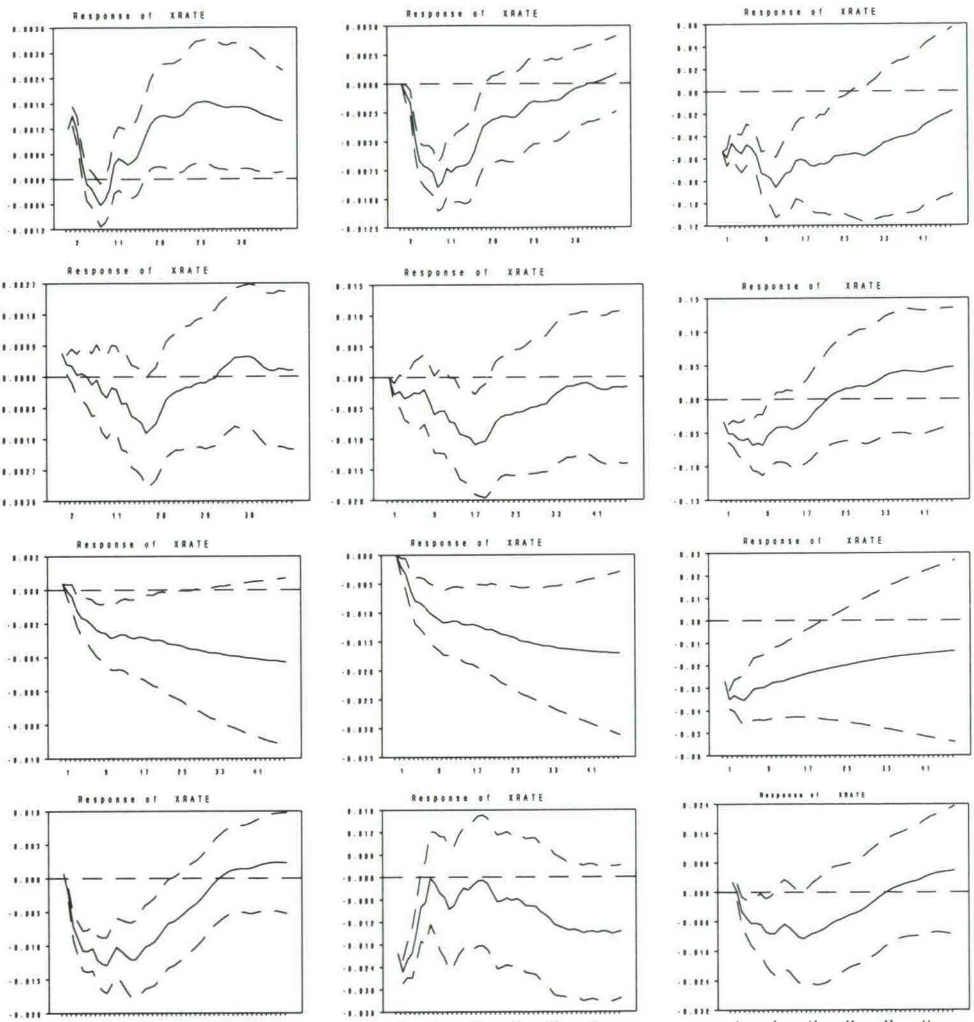
Fully Recursive Scheme | Semi-Recursive Scheme | Structural Model



Notes: The rows refer to the estimated responses of the nominal exchange rate to a one standard deviation foreign monetary policy shock obtained using Canadian, German, Japanese and British data respectively. The columns correspond to the various identification schemes used in estimating the responses.

Figure 5.5b. Monetary policy identification schemes and estimated exchange rate responses to respective foreign monetary policy shocks

Fully Recursive Scheme | Semi-Recursive Scheme | Structural Model



***Notes*:** The rows refer to the estimated responses of the nominal exchange rate to a one standard deviation foreign monetary policy shock obtained using Canadian, German, Japanese and British data respectively. The columns correspond to the various identification schemes used in estimating the responses.

Chapter 6

Comovements in Budget Deficits, Money, Interest Rates, Exchange Rates and the Current Account: Some Empirical Evidence.[1]

6.1 Introduction

During the 1980s the long-term interest rates in the US rose, the dollar appreciated and the current account balance deteriorated following a steady increase in the US budget deficit that began in 1981[2]. In Europe, since the unification of Germany, the movement of the deutsche mark (DM) and the long-term interest rates have been upwards due to the effects of the unification on the bond and foreign exchange markets. Branson(1993) argues that the current appreciation of the German DM in real terms is mainly due to fiscal expansion by the federal government following from the re-unification. Also the recent devaluation and the floating of the Swedish kronor took place at a time of relatively rising interest rates, a substantial and rising budget deficits and worsening current account balance situation. The simultaneous occurrences in the US economy, and the recent events in Sweden and Germany also suggest that there is some relationship among budget deficits, real exchange rate, interest rate differential and the current account balance. Indeed, the importance of budget deficits has been recognised and incorporated into many theoretical models, especially, Portfolio Balance Models (PBM) models. (See Branson (1983), (1984), (1985), (1988) and (1993) and Branson and Marcheese(1988)).

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[1] This chapter is a slightly revised version of Ibrahim and Kumah(1996), "Comovements in Budget Deficits, Money, Interest Rates, Exchange Rates and the Current Account: Some Empirical Evidence", ***Applied Economics***, 28, 117 - 130.

[2]. In the period 1981 - 1985, budgetary deficits in the US rose from almost zero to a total of $140 billion in 1985. In this same period, there was simultaneous appreciation of the US dollar in real and nominal terms (in comparison to the other major currencies of the world) as well as a deterioration in the current account balance from a current account surplus of $6.0 billion in 1981 to an annual rate of $120 billion in deficits by mid 1985 after a series of large and recurrent budget deficits. See Nakibullah(1993) for further dilation on the situation.

In spite of the apparent recognition in the literature on exchange rate determination of the importance of the effects of fiscal policy, especially, budget deficits on the long-term interest rate, the real exchange rate and hence the current account balance, not much empirical work has been done on the effects of budget deficits on the real exchange rate. This might be due partly to the limitations of the standard econometric approach that is mainly concerned with structural estimation. However, the recent developments in time series analysis, especially, vector auto-regression (VAR), cointegration, and common trends analysis have provided the means for analysing and establishing the short- and long-run relationships among various variables. Indeed, Branson(1983) states that in summary it appears that the VAR results in that chapter are quite consistent with the theory that attempts to integrate money, relative prices, and the current account balance, into a framework explaining movements in the real exchange rate.

Recent works that have provided the basis for this chapter include, Stockman(1983), Mussa (1986), and Baxter and Stockman(1989). Eichenbaum and Evans(1995) building on the works of the above authors employs a VAR approach to present new empirical evidence that expansionary monetary policy shocks generate substantial, persistent depreciations in the US nominal and real exchange rate. This chapter builds on Eichenbaum and Evans(1995) by incorporating expansionary fiscal policy shocks, especially, budget deficits into the analysis. In this chapter we shall adopt a VAR approach to analyze the effects of budget deficits on the real exchange rate, interest rate differential, and the current account balance within a framework of the portfolio balance model (PBM) of Branson(1993) and Branson and Marcheese(1988). The rest of the chapter is organised as follows. Section two presents the theoretical model. Section three deals with the analysis of the data and empirical results. The fourth section contains the summary and conclusions of the chapter.

6.2 The Theoretical Framework

There are a number of models that attempt to capture and explain the effects of fiscal as well as monetary innovations on the exchange rate (e_t), interest rate differential (r_t^D) - which is defined as the domestic interest rate minus the foreign interest rate, and the current account balance (*ca*). However, since we are more specifically interested in the effects of the structural deficit on the dependent variables mentioned above we adopt as our background model Branson's two-country real model of 1993. Most of the expositions and arguments raised in that paper can be found in other papers as well - see for instance Branson(1983, 1985 and 1988), Branson and Marcheese (1988), Branson and Love(1988), Feldstein(1986), and Nakibullah(1993). We choose the two-country real model not only for its simplicity and ability to capture explicitly the effects of the structural deficit on the three variables of interest - the interest rate differential, the exchange rate (real and nominal) and the current account balance - but also because it is suitable for the analysis of the variables of interest as mentioned above[3].

6.2.1 *The Basic Model*

In the two-country real model of Branson(1993) the financial variables are the long-term interest rates and the real exchange rate which are *jump* variables. The current account balance on the other hand is assumed to change the debt positions of the two countries gradually (this is to say it is a *sticky* variable). The model predicts that a fiscal or saving shock in say the first (home) country leads to an overshoot of the real exchange rate between the two countries, and also causes the real interest rate to overshoot its long-run level in the second (foreign) country whilst it undershoots its long-run level in the home country.

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[3]. The model assumes aggregate demand fixed/exogenous and hence, in our context, causally supersedes the budgetary deficit, the money supply and the three variables of interest. The model is therefore consistent with the Wold causal ordering of the data set that we adopt in this chapter. (See the next section for details on the data).

The basic relationships of the model are represented by the following IS equilibria for both the domestic and the foreign countries, an arbitrage condition and a relation describing the current account dynamics.

$$bdef_t = NS(r_t) - ca(e_t,B_t) \quad (6.1)$$
$$bdef_t^* = NS(r_t^*) + ca(e_t,B_t) \quad (6.2)$$
$$r_t^D = \hat{e}_t + \rho(B_t) \quad (6.3)$$
$$\dot{B}_t = -ca(e_t,B_t) \quad (6.4)$$

where

bdef denotes the home country's structural budgetary deficit,

NS denotes the home country's net-savings function of the private sector defined as the difference between private savings and investment expenditure, and is increasing in the interest rate, r_t,

ca_t denotes the current account balance; and it is an increasing function of the real exchange rate, e_t, and a decreasing function of the level of the domestic debt, B_t,

r^D_t denotes the interest rate differential between the two financial centers (home and foreign),

$\rho(B_t)$ denotes a risk premium as an increasing function of the domestic debt, B_t, and

$\hat{e}_t$ is the percentage change in the real exchange rate.

All starred variables are the foreign equivalents of their domestic counterparts. Under the assumption of full employment, NS and NS^* are the excesses of private savings over investment in the home and foreign country respectively. The real exchange rate is defined in terms of the home country's currency per unit of the foreign currency (so that an increase in e signifies a depreciation of the home currency). When the net debt of the home country to the foreign country is negative (i.e for $B < 0$) it implies that the home country is a creditor. The home country is a debtor when net debt to foreign is positive (i.e. for $B > 0$). Equation (6.3) is the arbitrage condition that links the financial markets. The risk premium is an increasing function of net debt, B.

6.2.2 Short-Run Dynamics and the Effects of Budgetary Deficits

The short-run dynamics of the model are captured in figure 6.1a below. The $\dot{B}_t = 0$ line is the locus of points along which $ca = 0$ whereas the $\hat{e} = 0$ line is the line along which the real exchange rate is not expected to change. Both lines have positive slopes but that of the latter line is greater than that of the former[4]. The positivity of the slopes is explained by the fact that along the $\dot{B}_t = 0$ line increases in the real exchange rate which increase the current account balance (in conformity with the Marshall-Lerner condition) require an increase in the domestic debt in order to leave the current account balance unchanged. For the $\hat{e} = 0$ line, the arbitrage relation (6.3) with $\hat{e} = 0$, implies that any increase in the risk premium requires an increase in the interest rate differential, r^D_t which is equivalent to an increase in r_t or a decrease in r^*_t. Suppose r_t increases (which in turn increases domestic net savings) , then given $bdef_t$ in equation (6.1), we require an increase in ca (which requires an increase in either e_t or B_t or both) - hence as B_t increases e_t increases and ca increases driving a wedge between the slopes of the two lines under consideration. For a detailed discussion of the dynamics as summarized in figure 6.1a below see Branson(1993).

[4]. The respective slopes can be explicitly derived. Notice that from equation (6.4) the slope of the first locus under consideration $[\delta e/\delta B] = -\ ca_B/ca_e > 0$ since $ca_B < 0$. However from equations (6.1) - (6.3) the slope of the second locus of points (for which $\hat{e} = 0$) is derived as $[\delta e/\delta B]$ [for $\hat{e}=0$ and $\delta bdef = \delta bdef^* = 0] = (\rho B - 2ca_B)/2ca_e > 0$. In the notation used above generally ca_χ is the partial derivative of the $ca(e_t, b_t)$ function with respect to χ.

Fig. 6.1 (a): The Dynamics of the Model

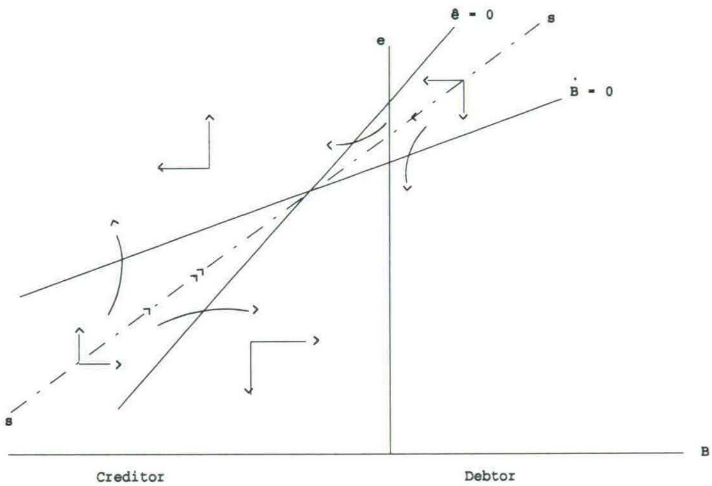
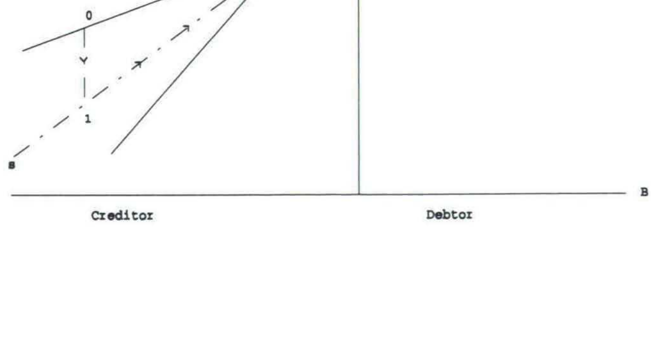


Fig. 6.1(b): Effects of Innovations (Unanticipated Shocks) in the Structural Deficit (Given that initially the $ca_t = 0$)



From fig. 6.1b on the previous page we see that given an initial creditor position/equilibrium at 0 a shock to the structural deficit (or unanticipated budget deficit increases) instantly appreciates the domestic currency but with time this translates into worsening current account situations (which show up as increases/reductions in debt/credit) and a depreciated currency. The exchange rate displays the overshooting phenomenon. These are some of the issues that we investigate empirically in the next section using VAR modelling.

6.3 VAR Modelling of the Open Economy

According to the Granger Representation Theorem (GRT) if a p x 1 vector X_t generated by $(I - L)X_t = \delta + C(L)\varepsilon_t$ is cointegrated (i.e. X_t is $CI(1,1)$) then there exists a Vector Autoregression (VAR), an Error Correction (EC) as well as a Moving Average (MA) representations of X_t (See Engle and Granger(1987). Hylleberg and Mizon (1989) have extended the GRT to include both a Bewley and a Common Trends representations. Our task in this section is to analyze the cointegration properties of the data set using the "Johansen-procedure" (see for instance Juselius and Johansen(1990)) and proceed to discuss the VAR representation. This section looks at only the theoretical aspects including derivations of the impulse-response functions as well as the forecast error decompositions if the cointegration rank is different from zero.

6.3.1 Preliminary Data Analysis

The data utilized are seasonally unadjusted quarterly data collected from several issues of *International Financial Statistics*. To conserve space we have not included time series plots of the data in this chapter. In order to compare the relative effects of money and the budgetary deficit on the interest rate differential, the real exchange rate and the current account balance we augment the variables of our model with the set [*log gnp, log m*]. Thus the set of variables (X_t) analyzed in this chapter has as its elements - in the Wold ordering - the logarithm of *gnp* (*log gnp*), the structural budgetary deficit (*bdef*), the logarithm of broad money (*log m*), the interest rate differential (r^D), the real exchange rate (*e*), and the current account balance (*ca*) for Germany, Japan, Sweden and the United Kingdom over the period 1974:1 - 1992:4. The interest rate differential in all cases is the difference between the domestic three month treasury bill discount rate and the Federal Funds Rate of the United States. The real exchange rate is calculated as the domestic currency price of one US dollar multiplied by the ratio of the US CPI to the domestic CPI. Thus defined this way, an increase in e denotes a depreciation of the domestic currency. We investigate the time series characteristics of the data by utilizing the multivariate error correction model of Johansen - the Johansen procedure as outlined below.

6.3.2 The Johansen Procedure

The existence of long-run equilibrium (stationary) relationships among economic variables is referred to in the literature as cointegration. A set of variables, X_t is said to be cointegrated of order (*d,b*) - denoted *CI(d,b)* - if X_t is integrated of order *d* and there exists a vector, β, such that $\beta' X_t$ is integrated of order (*d - b*)[5]. The most common test for cointegration is the two-step procedure of Engle and Granger(1987) which performs the tests in a univariate setup. Recent developments in the literature include the Johansen procedure (see Johansen(1988) and Johansen and Juselius(1990)). These studies examine the question of cointegration and provide not only an estimation methodology but also explicit procedures for testing for the number of cointegrating vectors as well as for restrictions suggested by economic theory - in a multivariate setting.

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[5]. A variable is said to be integrated of order *z* - denoted *I(z)* if the said variable becomes covariance stationary after differencing z times. See Cryer(1986) for a further definition of the concept of stationarity

Following Johansen and Juselius(1990) let the p variables under scrutiny follow a vector autoregression of order k as below

$$X_t = \mu + \Pi_1 X_{t-1} + \ldots + \Pi_k X_{t-k} + \epsilon_t \qquad (6.5)$$

where $\varepsilon_1, \ldots \varepsilon_T$ are the innovations of this process and are assumed to be drawn from a p-dimensional i.i.d. Gaussian distribution with covariance Ξ and $X_{-k+1}, \ldots, X_0$ are fixed. Letting Δ represent the first difference operator (where $\Delta = (1 - L)$ for L being the lag operator such that $L^i Z_t = Z_{t-i}$), equation (6.5) can be reparameterised into the equivalent form - necessary for isolating the crucial impact matrix, $\Pi(1)$ - as presented below

$$\Delta X_t = \mu + \Gamma_1 \Delta X_{t-1} \ldots + \Gamma_{k-1} \Delta X_{t-k+1} - \Pi X_{t-k} + \psi D_t + \epsilon_t \qquad (6.6)$$

where

$$\Gamma_i = -I + \sum_{j=1}^{i} \Pi_j \qquad (\textit{for } i = 1, \ldots, k-1), \qquad \Pi = I + \sum_{1}^{k} \Pi_j ,$$

ε_t is i.i.d normal with a zero mean vector and covariance Ξ, and D_t denotes centered (that is to say their sum is zero) seasonal dummies. The rank, r, of the long run matrix, Π, defines the number of cointegrating vectors/relationships in the data. When $0 <$ rank (Π) $= r - p$, then Π can be written as $\Pi = \alpha\beta'$ where β is interpreted as a $p \times r$ matrix of cointegrating vectors and α as a $p \times r$ matrix of error correcting parameters.

Johansen and Juselius(1990) derive the cointegrating vector, β, by solving for the eigenvalues of

$$| \lambda S_{kk} - S_{k0} S_{00}^{-1} S_{0k} | = 0$$

where S_{00} is the moment matrix of the residuals from ordinary least squares regression of ΔX_t on $\Delta X_{t-1}, \ldots, \Delta X_{t-k+1}$, S_{kk} is the residual moment matrix from ordinary least squares regression of ΔX_{t-k} on ΔX_{t-k+1}, and S_{0k} is the cross-product moment matrix. The cointegrating vector, β, is solved out as the eigenvectors associated with the r largest, statistically significant eigenvalues derived above (and ordered in descending order) using two test statistics - the 'maximum eigenvalue statistics' and the 'trace statistics'. The first statistics tests the hypothesis that there are $r = s$ (where $s \leq p$) cointegrating vectors (against the alternative of $r = s + 1$) by calculating the maximum likelihood test statistics - LR_{max} - as below

$$LR_{max} = - T \ln(1 - \lambda_{s+1})$$

where T is the sample size and λ_{s+1} is an estimated eigenvalue. The second statistics, the 'trace statistics' (LR_{trace}) tests the hypothesis that there exists at most r cointegrating vectors by calculating the likelihood test statistics given by

$$LR_{trace} = - T \sum_{r+1}^{p} \ln(1 - \lambda_i)$$

where T, as before, is the sample size and $\lambda_{r+1}, \ldots, \lambda_p$ are the estimated $p - r$ smallest eigenvalues. The distributions of these test statistics as tabulated in Osterwald-Lenum(1992) are non-standard and depend only on the number of degrees of freedom, $p - r$.

To derive the lag length to be used in this procedure we performed a likelihood ratio test (the results of which are not included here) which, at the 5% significance level, pointed to a lag length of four as optimal for all countries. The same lag length is supported by the multivariate performance of the residuals as against those for other lags (eg. for lag six). See the results of the cointegration tests as presented in table 6.1 below.

Table 6.1: Cointegration Tests for the UVAR(4) representation for $X_t = [log.\ gnp\ log\ m\ r^D\ e\ ca]'$

a) Germany

	Max. Eigen Value $-T\ \ln(1-\lambda_i)$			Trace Statistics $-T \sum \ln(1-\lambda_i)$		
H_0	**H_1**	**LR_{max}**	**P-Value**	**H_1**	**LR_{trace}**	**P-Value**
r = 0	r ≥ 1	54.94	< 1%	r = 1	132.79	< 1%
r ≤ 1	r ≥ 2	28.38	≈ 20%	r = 2	77.85	< 1%
r ≤ 2	r ≥ 3	25.49	> 5 %	r = 3	49.47	> 5%
r ≤ 3	r ≥ 4	16.11	> 20%	r = 4	23.98	≈ 20%
r ≤ 4	r ≥ 5	7.65	> 20%	r = 5	7.87	> 20%
r ≤ 5	r = 6	0.22	> 20%	r = 6	0.22	> 20%

Notes: *The P-Values (indicating the degree of credibility assigned the null hypothesis by the data) are taken from Osterwald-Lenum (1992) table 1*.*

Multivariate and Univariate Error Diagnostics (Given that r = 2).

(i) ***Multivariate Tests for autocorrelation.***

B-P(18) = 0.9165;
LM(1) = 3.6154, p-value = 1.0000;
LM(4) = 27.035, p-value = 0.8598

(ii) ***Univariate tests.***

Error	B-P. Q(18)/14	ARCH(4)	J-B. NORM.
ε_1	0.786	0.378	766.617
ε_2	0.481	27.047	2.519
ε_3	0.439	0.152	2220.646
ε_4	1.033	12.908	4.328
ε_5	0.987	7.669	15.937
ε_6	0.697	5.597	0.175

Notes: *Under subsection (i) B-P(18) is the multivariate version of the Box-Pierce test for autocorrelation based on 18 autocorrelations. It has an asymptotic χ^2 ($n^2(k-p)$) - distribution, where k denotes the number of autocorrelations and p is the lag length. LM(1) and LM(4) are the lagrangian multiplier test for autocorrelation at lags 1 and four respectively. Under subsection (ii) we have the univariate B-P(18) test for autocorrelation with a limiting $\chi^2(k-p)$ distribution. ARCH(4) is a test for fourth order autoregressive conditional heteroscedasticity and J-B. Norm is the Jacque-Bera test for normality - both tests have a $\chi^2(4)$ distribution.*

b) **Japan**

	Max. Eigen Value $-T \ln(1-\lambda_i)$			Trace Statistics $-T \sum \ln(1-\lambda_i)$		
H_0	H_1	LR_{max}	P-Value	H_1	LR_{trace}	P-Value
$r = 0$	$r \geq 1$	67.679	< 1%	$r = 1$	161.66	< 1%
$r \leq 1$	$r \geq 2$	35.104	≈ 2.5%	$r = 2$	93.983	< 1%
$r \leq 2$	$r \geq 3$	29.501	< 2.5 %	$r = 3$	58.880	< 1%
$r \leq 3$	$r \geq 4$	17.911	> 10%	$r = 4$	29.379	≈ 5%
$r \leq 4$	$r \geq 5$	9.579	> 20%	$r = 5$	11.468	> 10%
$r \leq 5$	$r = 6$	1.889	> 10%	$r = 6$	1.889	> 10%

Notes: *The P-Values (indicating the degree of credibility assigned the null hypothesis by the data) are taken from Osterwald-Lenum (1992) table 1*.*

Multivariate and Univariate Error Diagnostics (Given that r = 3).

(i) ***Multivariate Tests for autocorrelation.***

B-P(18) = 1.0972;
LM(1) = 8.5366, p-value = 1.0000;
LM(4) = 38.948, p-value = 0.3385

(ii) ***Univariate tests.***

Error	B-P. Q(18)/14	ARCH(4)	J-B. NORM.
ε_1	0.641	2.088	2.618
ε_2	1.478	0.396	906.797
ε_3	0.805	2.013	5.128
ε_4	0.834	3.022	85.958
ε_5	0.992	3.676	1.292
ε_6	1.566	1.916	11.620

Notes: *Under subsection (i) B-P(18) is the multivariate version of the Box-Pierce test for autocorrelation based on 18 autocorrelations. It has an asymptotic χ^2 $(n^2(k-p))$ - distribution, where k denotes the number of autocorrelations and p is the lag length. LM(1) and LM(4) are the lagrangian multiplier test for autocorrelation at lags 1 and four respectively. Under subsection (ii) we have the univariate B-P(18) test for autocorrelation with a limiting $\chi^2(k-p)$ distribution. ARCH(4) is a test for fourth order autoregressive conditional heteroscedasticity and J-B. Norm is the Jacque-Bera test for normality - both tests have a $\chi^2(4)$ distribution.*

c) **Sweden**

	Max. Eigen Value $-T \ln(1-\lambda_i)$			Trace Statistics $-T \sum \ln(1-\lambda_i)$		
H_0	H_1	LR_{max}	P-Value	H_1	LR_{trace}	P-Value
$r = 0$	$r \geq 1$	53.534	< 1%	$r = 1$	159.22	< 1%
$r \leq 1$	$r \geq 2$	40.573	< 1%	$r = 2$	105.69	< 1%
$r \leq 2$	$r \geq 3$	30.746	< 2.5 %	$r = 3$	65.121	< 1%
$r \leq 3$	$r \geq 4$	16.738	> 10%	$r = 4$	34.376	< 2.5%
$r \leq 4$	$r \geq 5$	10.708	> 10%	$r = 5$	17.637	< 2.5%
$r \leq 5$	$r = 6$	6.929	> 1%	$r = 6$	6.929	< 1%

Notes: *The P-Values (indicating the degree of credibility assigned the null hypothesis by the data) are taken from Osterwald-Lenum (1992) table 1*.*

Multivariate and Univariate Error Diagnostics (Given that r = 3).

(i) ***Multivariate Tests for autocorrelation.***

B-P(18) = 0.9508;
LM(1) = 7.7446, p-value = 1.0000;
LM(4) = 23.531, p-value = 0.9456

(ii) ***Univariate tests.***

Error	B-P. Q(18)/14	ARCH(4)	J-B. NORM.
ε_1	0.612	2.924	407.327
ε_2	0.720	4.817	44.891
ε_3	0.645	20.164	45.879
ε_4	0.763	3.356	1.472
ε_5	0.892	2.042	54.816
ε_6	1.472	0.547	185.617

Notes: *Under subsection (i) B-P(18) is the multivariate version of the Box-Pierce test for autocorrelation based on 17 autocorrelations. It has an asymptotic χ^2 $(n^2(k-p))$ - distribution, where k denotes the number of autocorrelations and p is the lag length. LM(1) and LM(4) are the lagrangian multiplier test for autocorrelation at lags 1 and four respectively. Under subsection (ii) we have the univariate B-P(18) test for autocorrelation with a limiting $\chi^2(k-p)$ distribution. ARCH(4) is a test for fourth order autoregressive conditional heteroscedasticity and J-B. Norm is the Jacque-Bera test for normality - both tests have a $\chi^2(4)$ distribution.*

d) UK

	Max. Eigen Value $-T\ln(1-\lambda_i)$			Trace Statistics $-T\sum\ln(1-\lambda_i)$		
H_0	H_1	LR_{max}	P-Value	H_1	LR_{trace}	P-Value
$r = 0$	$r \geq 1$	37.373	> 5%	$r = 1$	108.85	< 1%
$r \leq 1$	$r \geq 2$	29.819	> 10%	$r = 2$	71.486	≈ 2.5%
$r \leq 2$	$r \geq 3$	17.155	> 20 %	$r = 3$	41.667	> 10%
$r \leq 3$	$r \geq 4$	13.560	> 20%	$r = 4$	24.513	> 10%
$r \leq 4$	$r \geq 5$	9.568	>20%	$r = 5$	10.953	> 20%
$r \leq 5$	$r = 6$	1.384	> 20%	$r = 6$	1.384	> 20%

Notes: *The P-Values (indicating the degree of credibility assigned the null hypothesis by the data) are taken from Osterwald-Lenum (1992) table 1*.*

Multivariate and Univariate Error Diagnostics (Given that r = 2).

(i) ***Multivariate Tests for autocorrelation.***

B-P(17) = 1.0539;
LM(1) = 6.8804, p-value = 1.0000;
LM(4) = 33.540, p-value = 0.5861

(ii) Univariate tests.

Error	B-P. Q(17)/13	ARCH(4)	J-B. NORM.
ε_1	0.677	0.198	1126.995
ε_2	1.167	0.560	8.007
ε_3	0.918	2.736	91.504
ε_4	1.681	3.011	3.521
ε_5	0.618	0.395	0.395
ε_6	1.176	4.380	4.380

Notes: *Under subsection (i) B-P(17) is the multivariate version of the Box-Pierce test for autocorrelation based on 17 autocorrelations. It has an asymptotic χ^2 $(n^2(k-p))$ - distribution, where k denotes the number of autocorrelations and p is the lag length. LM(1) and LM(4) are the lagrangian multiplier test for autocorrelation at lags 1 and four respectively. Under subsection (ii) we have the univariate B-P(17) test for autocorrelation with a limiting $\chi^2(k-p)$ distribution. ARCH(4) is a test for fourth order autoregressive conditional heteroscedasticity and J-B. Norm is the Jacque-Bera test for normality - both tests have a $\chi^2(4)$ distribution.*

Table 6.1 (contd).

A. Likelihood Ratio Tests for Staionarity

	ln gnp	*bdef*	*ln m*	r^D	*e*	*ca*
Germany	20.78	6.32	22.26	21.42	12.40	9.97
Japan	19.29	11.50	20.61	15.18	17.20	25.67
Sweden	3.63	17.76	10.35	7.94	7.71	10.61
UK	26.51	20.95	24.99	13.07	7.29	25.69

Notes: *The likelihood ratio statistics has a $\chi^2(n\text{-}r)$ distribution where n is the number of elements in the X_t vector (n = 6 for all countries considered) and r denotes the cointegration rank in each case. Therefore the critical 5% significance levels are 9.49, 7.81, 7.81 and 9.49 for Germany, Japan, Sweden and the UK respectively.*

B. Likelihood Ratio Tests for Exclusion

	ln gnp	*bdef*	*ln m*	r^D	*e*	*ca*
Germany	21.70	2.63	16.09	7.92	26.70	19.99
Japan	38.16	23.34	35.98	38.98	15.28	18.26
Sweden	14.14	5.77	1.57	21.81	14.15	23.98
UK	11.74	0.53	5.99	10.34	10.40	16.18

Notes: *The likelihood ratio statistics has a $\chi^2(r)$ distribution where r denotes the cointegration rank. Therefore the critical 5% significance levels are 5.99, 7.81, 7.81 and 5.99 for Germany, Japan, Sweden and the UK respectively.*

The results support cointegration ranks of 2, 3, 3 and 2 for Germany, Japan, Sweden and the UK respectively. Likelihood ratio tests for stationarity (given the respective cointegration ranks) revealed *bdef* in the case of Germany, *log. gnp* and the exchange rate in the case of Sweden, and exchange rate in the case of the UK to be integrated of order zero - *I(0)*. All other variables are *I(1)* for all the countries. The likelihood ratio test for exclusion of variables from the X_t vector suggested that the *bdef* be excluded for Germany, Sweden and the UK, and *log. m*, as well, be excluded for Sweden. However considering that our main concern is about the effects of budgetary deficits on r^D, *e* and *ca* we decided to include the *bdef* nevertheless. A similar analysis is carried out for the X_t vector with the nominal exchange rate (*s*) substituted for the real exchange rate (e). The results for this second analysis are not reported here since they do not differ significantly from those presented for the real exchange rate.

From the results presented above and summarized in table 6.1 we infer that the variables of interest are cointegrated and hence we proceed to utilize the VAR representation in our further analyses. Next, we present a summary description of the VAR representation and the derivation of the impulse-response functions and the forecast-error decompositions.

6.3.3 The VAR Representation

Since X_t is a p x 1 cointegrated stochastic vector, its k th-order autoregression can be expressed as

$$X_t = \alpha + \sum_{j=1}^{k} A_j X_{t-j} + \epsilon_t \qquad (6.7)$$

where ε_t is an p×1 vector of disturbances and α is an $p\times 1$ vector of constants. The $A_{j's}$ are $p\times p$ matrices that are, under some regularity conditions, determined by the orthogonality conditions

$$E[\varepsilon_t] = 0 \qquad \text{and} \qquad E[\varepsilon_t \mid X'_{t-j}] = 0_{p\,x\,p}, \quad j = 1,2 \ldots, k.$$

The ε_t process is termed the process of innovations - the process of one-step-ahead prediction errors - since it is that part of X_t that cannot be predicted from the k lagged X_t's. If k is assumed to be large enough, as we assume below, then the ε_t vector is serially uncorrelated- i.e. $E[\varepsilon_t \varepsilon'_{t-s}] = 0$ for $s \neq 0$ - but contemporaneously correlated such that $E[\varepsilon_t \varepsilon'_t] = \Xi$ which is in general not a diagonal matrix.

Solving equation (6.7) for X_t backward in terms of the innovations we get the vector moving average representation

$$X_t = \alpha' + \sum_{j=0}^{\infty} C_j \epsilon_{t-j} \tag{6.8}$$

where α' is an $p\times 1$ vector of constants and C_j is an $p\times p$ matrix with $C_0 = I_p$. To construct a moving average representation with a disturbance process that is orthogonal contemporaneously as well as at all lags, consider the transformation $\varepsilon_t = Fu_t$ where F is an $p\times$p lower triangular matrix such that (due to the least-squares orthogonality condition $E[u_{it}\varepsilon_{jt}] = 0$, for $i,j = 1,2,...,p$ and $i \neq j$) the element $f_{ij} = E[\varepsilon_{it}\varepsilon_{jt}] / E[\varepsilon^2_{jt}]$ for $i, j = 1,2,...,p$ and $i \neq j$. By construction u_{it} and u_{jt} (for $i, j = 1,2,...,p$ and $i \neq$ j) are orthogonal and therefore a moving average representation in terms of orthogonal innovations at all lags is given by

$$\begin{aligned} X_t &= \alpha' + \sum_{j=0}^{\infty} C_j\, F\, u_{t-j} \\ &= \alpha' + \sum_{j=0}^{\infty} D_j\, u_{t-j} \end{aligned} \tag{6.9}$$

where $D_j = C_jF$. This is the Choleski decomposition of the u_t vector. One could also use the structural decomposition suggested by Bernanke. (See Doan(1990)). Hence generally any choice of $u_t = F^{1}\varepsilon_t$ that makes $E[u_tu'_t] = F^{1}\ \Xi\ F^{\prime 1}$ diagonal can be utilized to derive a moving average representation in terms of an innovation process that is orthogonal contemporaneously and at lags and leads. Equation (6.9) can then be used to derive the impulse-responses.

From equations (6.8) and (6.9) the k-step-ahead error in forecast X_t from its own past is given by

$$\begin{aligned} X_t - E_{t-k}X_t &= C_0\epsilon_t + \dots + C_{k-1}\epsilon_{t-k+1} \\ &.= D_0u_t + \dots + D_{k-1}u_{t-k+1} \end{aligned} \tag{6.10}$$

where $E_{t-k}\, X_t$ is the least-squares forecast of X_t given X_{t-k}, X_{t-k-1}, ...

The variance-covariance matrix of the k-step-ahead prediction errors is given by the expression

$$E\,[\,(\,X_t - E_{t-k}X_t\,)(\,X_t - E_{t-k}X_t\,)'\,] = D_0\,E[u_t u_t']D_0' \;+\; \dots + D_{k-1}\,E[u_t u_t']D_{k-1}' \,. \qquad (6.11)$$

The diagonal terms of the above expression gives the decomposition of the variance of the k-step-ahead error into the parts attributable to shocks in the n components of u_t. The procedure described above is utilized in analysing the six-variable vector X_t for each of the four countries mentioned above using RATS[6]. The empirical results are presented in the next section.

6.4 Empirical Results

The empirical results that we have obtained from this investigation of the open economy using the VAR approach are rather complicated and difficult to interpret. However we discuss some interpretations based on the theoretical model presented in section 6.2 above as well as on general international monetary theory. The interpretations are as presented under the various subheadings below. Due to the unclarity of the results we shall limit ourselves to the first - four quarters only (that is to say only very short-run effects of fiscal and monetary shocks on the endogenous variables are considered). Figures 6.2 & 6.3 and table 6.2 contain the main empirical results that we discuss below. Notwithstanding this, figures 6.4 and 6.5 are also included here to provide additional information to readers interested in the effects of i) shocks to the interest rate differential (as in figure 6.4); and ii) real exchange rate innovations (as in figure 6.5)

[6]. See Doan (1990) for a description of the algorithms used.

6.4.1 Effects of Budgetary Deficits

According to the theoretical model used in this study, an increase in the structural budget deficit causes the real exchange rate to jump down (i.e. a sudden appreciation of the domestic currency) creating a current account deficit in the home country *(in the short run)*. This will require an increase in the domestic interest rate and/or a fall - or a less than proportionate increase - in the foreign interest rate as required in equations (6.1) and (6.2). The rise in interest differential must be equal to the sum of the new expected rate of depreciation of the domestic currency and the risk premium - as in equation (6.3).

From the impulse responses presented in figure 6.2, first, we observe that a one standard deviation shock to the domestic budget deficit (hereafter referred to as fiscal shock) do increase the interest rate differential given the Japanese and UK data over the first-four quarters. In the case of Sweden this positive effect is felt during the fifth-sixth quarter whereas in Germany the data does not seem to succeed in capturing the effects of fiscal shocks on the interest rate differential - this may be attributed to inaccuracies in the data set and/or misspecification of the basic model utilized.

Secondly, we observe that increases in the structural budget deficit do appreciate the domestic real exchange rate at least in the first-three quarters for all the countries considered in the study. Our finding here is that the econometric results do confirm the prediction of the theory that fiscal shocks worsen the current account balance even though the effects do not appear to be statistically significant.

6.4.2 Effects of Monetary shocks

Some of the main predictions of the International monetary theory and policy is that domestic monetary shocks do increase the interest rate differential (i.e. $r_t^D = r_t - r_t^*$), depreciate the real exchange rate and improve the current account balance. However, from the impulse-responses presented in figure 6.3, only aspects of this theoretical prediction can be traced from the data set used for all countries considered. We find that for the Japanese and Swedish data, monetary shocks significantly reduce the interest rate differential at least over the first-three quarters but have no significant effect on the real exchange rate but worsens the current account balance significantly (especially in the case of Japan and the UK).

On the whole the empirical results derived from the impulse- responses are rather unsatisfactory and hence could not yield very conclusive evidence as to the relative performance of fiscal and monetary shocks in terms of their respective effects on the interest rate differential (r_t^D), the real exchange (e_t) and the current account balance (ca_t). We therefore proceed to analyze the k-step forecast error decompositions (FEDs) as presented in table 6.2.

6.4.3 Relative Explanatory Power of Fiscal and Monetary shocks

The general impressions that we gather from an examination of the k-step forecast error decomposition (FEDs) presented in table 6.2 are summarized below (all the interpretations are limited to the first-four steps ahead forecast errors):

i) Compared with fiscal shocks monetary shocks explain a larger proportion of the forecast error variances of the interest rate differential (r_t^D) for all countries considered. For instance whereas innovations in the money supply explain between (0.02% and 1.70%, 8.39% and 9.09%, 28.11% and 38.55%, 5.32% and 5.33%) of the first-four steps ahead forecast errors for Germany, Japan, Sweden and the UK respectively, the corresponding figures for fiscal innovations are between (0.44% and 0.45%, 2.87% and 2.63%, 0.89% and 1.12%, 7.69% and 8.06%). A similar tendency is depicted in the case of the first-four steps ahead forecast error variances for the current accounts depicted by the corresponding percentages reported in table 6.2 for the respective countries considered.

ii) Fiscal innovations explain a relatively larger proportion of the first-four steps ahead forecast errors variances of the real exchange rate as compared to monetary innovations. Referring to table 6.2 again, we infer that fiscal innovations explain between (2.53% and 8.45%, 0.03% and 4.62%, 4.72% and 7.90%, 0.08% and 0.13%) whereas monetary innovations explain (0.02% and 0.47%, 0.01% and 1.26%, 0.45% and 4.52%, 0.14% and 10.60%) of forecast errors variances of the real exchange rate (e_t) for Germany, Japan, Sweden and the UK respectively.

In summary fiscal innovations performed relatively better than monetary innovations in explaining forecast error variances in the real exchange rate. This could be traced to the nature of foreign exchange market - more specifically, its responsiveness to signals form policy-making. It seems the real exchange rate responds more to fiscal innovations than to monetary innovations because issues relating to credibility as well as those bordering on the efficiency with which policy-makers control the economy, are inextricably intertwined with the process of expectation formation in the foreign exchange market. We conclude therefore that one of the main factors that influence the risk premium associated with domestic interest rate (as appropriately expressed in equation (6.3) in section 6.2) is the budget deficit (see Feldstein(1986) for a further dilation on this issue).

6.5 Summary and Conclusion

It is suggested that a change (i.e. an increase or a decrease) in the structural budget deficit has some potent effects on the interest rate differential, the real exchange rate and hence the current account balance. Even though theoretical advances have been made in the study of the effects of the structural budget deficit on the interest rate differential, real and nominal exchange rates as well as on the current account balance, not much empirical studies have been done.

This chapter is an attempt at an empirical analyses of the effects of fiscal and monetary innovations on the interest rate differential between financial centers, the real exchange rate and the current account balance using the familiar VAR approach. Our interest in the topic was aroused by the fact that most applications of cointegration and of VAR approach in the foreign exchange market do not include the budget deficit (or fiscal innovations) as one of the variables. (See for instance Eichenbaum and Evans(1995), Branson(1983), Hansen(1993) and Papell(1993)). We consider this a serious omission (even if the authors claim they are mainly interested in one particular theory of exchange rate determination) since the exchange rate responds more to fiscal innovations (as at least our findings show) than to monetary innovations.

Our empirical results indicate that variations in the interest rate differential between financial centers as well as of the current account balances of countries are explained more by monetary innovations than by fiscal innovations. Fluctuations in the real exchange rate seem to be better explained by fiscal innovations. This is however not to say that monetary innovations do not have anything to say in this respect. What our results indicate is that the policy-making provides actors in the foreign exchange market with certain signals crucial in expectations formation, and that the signals from fiscal innovations carry more weight than those from monetary innovations as far as real exchange rate fluctuations are concerned. This result gives evidence to the importance of including fiscal innovations in any cointegration and VAR analyses of exchange rate fluctuations. The results from our estimation of the impulse-responses are, however, not very significant and difficult to interpret. These may be due to the fact the effects of the fiscal and monetary shocks are not strong enough to be captured by our data set. It may also be due to the inadequacies in our data set or perhaps in our choice of model. Consequently, we realize the need for more empirical work in this area of interest to both policymakers and researchers.

6.6 References

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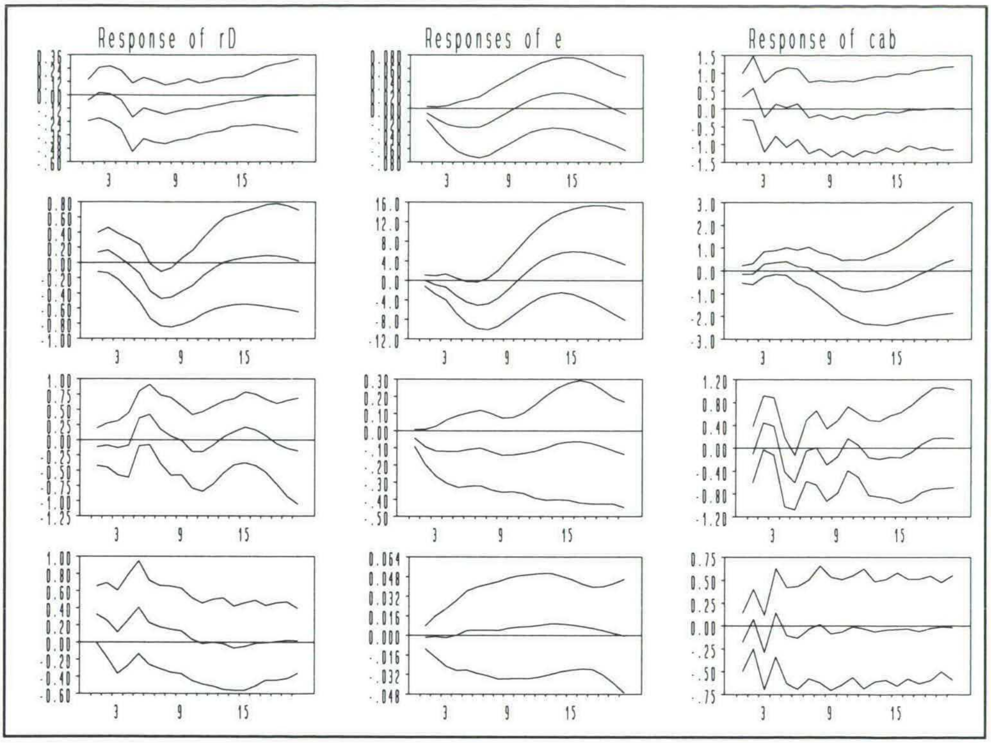
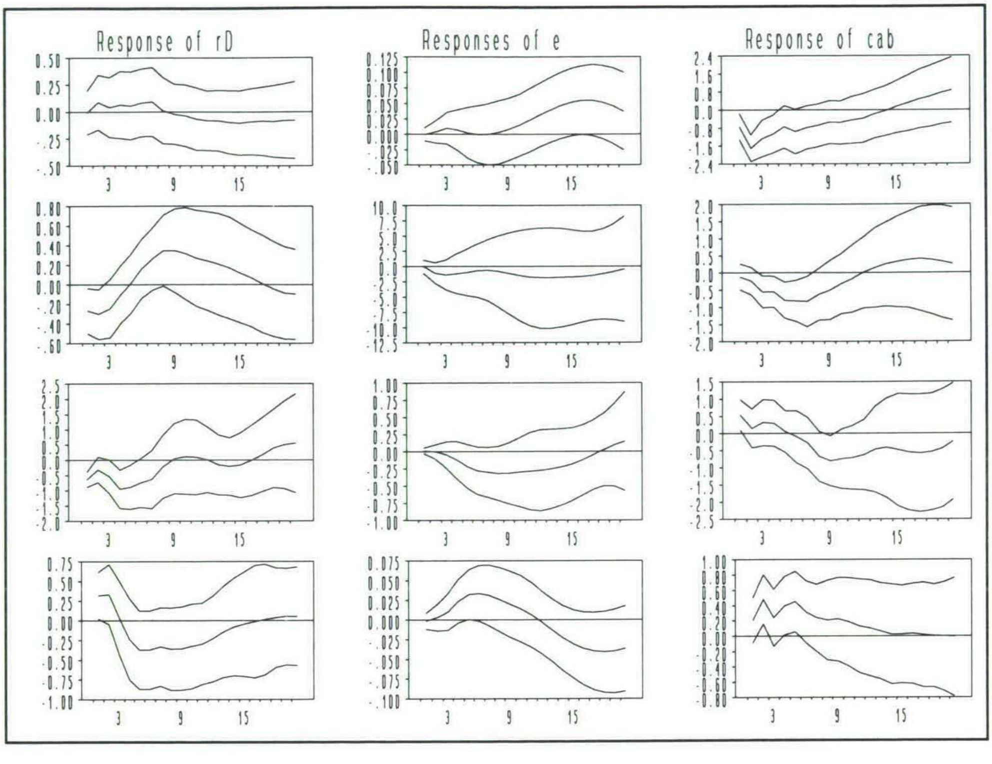
Figure 6.2: Responses of the interest rate differential (r^D), the real exchange rate (e), and the current account balance (ca) to a one standard deviation shock in the budgetary deficit ($bdef$)



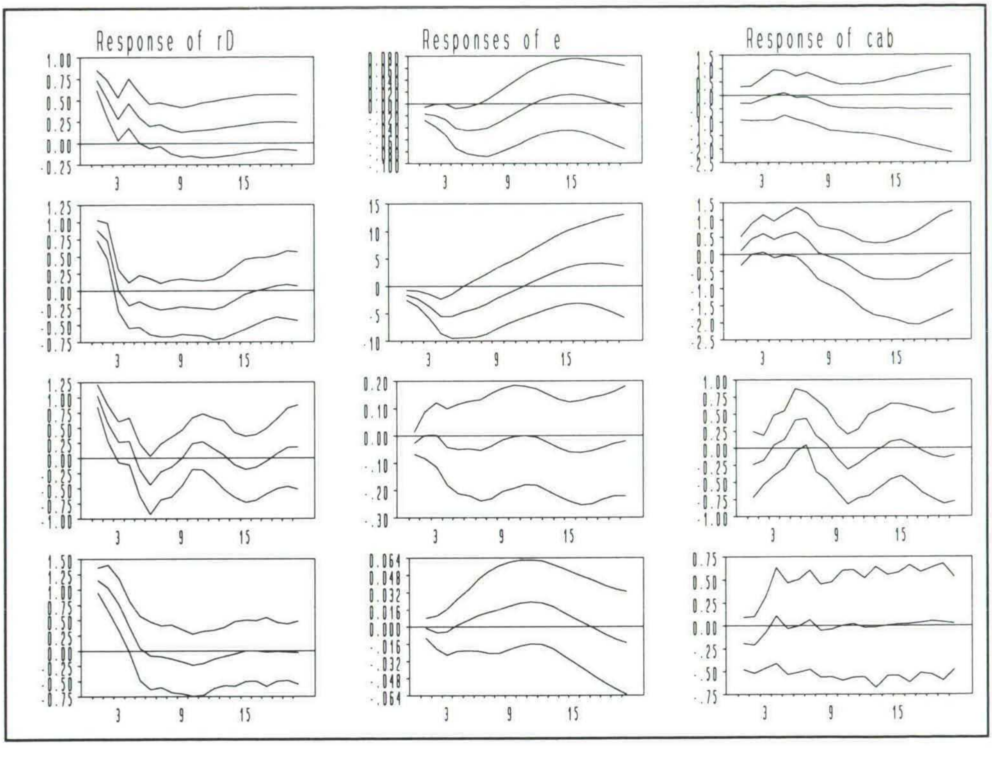
Notes: *The rows refer to impulse-response functions for Germany, Japan, Sweden and the U.K. respectively.*

Figure 6.3: **Responses of the interest rate differential (r^D), the real exchange rate (e), and the current account balance (ca) to a one standard deviation shock in the log of money ($ln\ m_t$)**



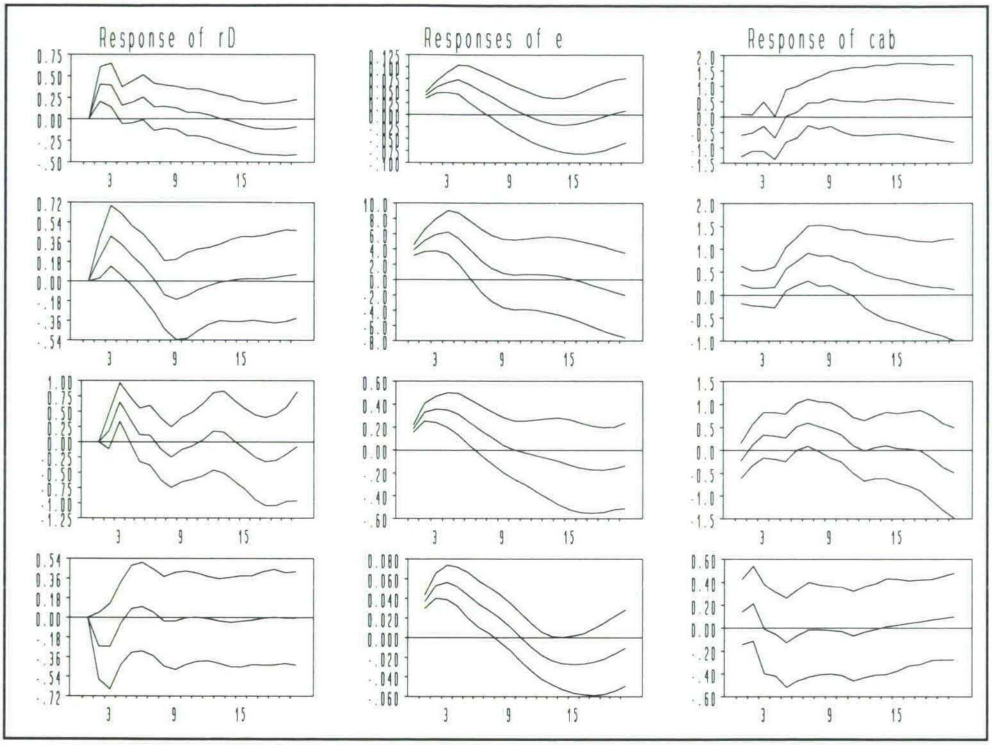
Notes: *The rows refer to impulse-response functions for Germany, Japan, Sweden and the U.K. respectively.*

Figure 6.4: **Responses of the interest rate differential (r^D), the real exchange rate (e), and the current account balance (ca) to a one standard deviation shock in the interest rate differential (r^D_t)**



Notes: The rows refer to impulse-response functions for Germany, Japan, Sweden and the U.K. respectively.

Figure 6.5: Responses of the interest rate differential (r^D), the real exchange rate (e), and the current account balance (ca) to a one standard deviation shock in the real exchange rate (e)



Notes: *The rows refer to impulse-response functions for Germany, Japan, Sweden and the U.K. respectively.*

TABLE 6.2: **Percentage of *k*-step-ahead Forecast Error Variances of the Interest Rate Differential (r^D), the Real Exchange Rate (*e*) and the Current Account Balance (*ca*) accounted for by shocks to *log. GNP*, Budgetary Deficit (*BDEF*) and *log. M***

A) GERMANY

k	VARIANCE OF k-STEP-AHEAD FORECAST ERROR	% VARIANCE IN *k*-STEP-AHEAD FORECAST ERROR EXPLAINED BY ORTHOGONALIZED INNOVATIONS IN:		
		log. GNP	*BDEF*	*log. M*
		Interest Rate Differential (r^D)		
1	0.5021	0.6698	0.4399	0.0208
2	0.9114	0.7116	0.2435	1.1115
3	1.1636	2.0211	0.1921	1.2184
4	1.3966	2.2233	0.4525	1.7033
5	1.5622	2.5455	3.2569	1.9788
10	2.0195	7.7257	6.3245	2.6134
15	2.3268	13.0738	6.5476	2.6054
20	2.8271	19.8736	5.4124	2.4041
		Real Exchange Rate (*e*)		
1	0.0002	0.6568	2.5310	0.0155
2	0.0006	6.9799	5.0120	0.1592
3	0.0129	11.8418	7.3883	0.5782
4	0.0217	13.9229	8.4458	0.4677
5	0.0298	16.7214	9.0139	0.3418
10	0.0548	35.9299	7.4818	0.5039
15	0.0779	36.3088	7.5872	9.9711
20	0.0960	30.3555	6.7306	19.9601
		Current Account Balance (*ca*)		
1	6.4633	0.2443	2.3004	8.3471
2	10.8050	4.1988	4.7003	31.3403
3	13.6245	3.5068	3.9992	37.4242
4	15.8770	6.3799	3.7021	39.5312
5	17.6862	8.4121	3.3241	38.4436
10	24.6641	10.7114	3.3359	35.9505
15	32.4353	20.9497	3.1353	28.5119
20	43.9927	30.7866	2.4772	25.0821

TABLE 6.2(contd): **Percentage of k-step-ahead Forecast Error Variances of the Interest Rate Differential (r^D), the Real Exchange Rate (e) and the Current Account Balance (ca) accounted for by shocks to *log. GNP*, Budgetary Deficit (*BDEF*) and *log. M***

B) JAPAN

k	VARIANCE OF k-STEP-AHEAD FORECAST ERROR	% VARIANCE IN k-STEP-AHEAD FORECAST ERROR EXPLAINED BY ORTHOGONALIZED INNOVATIONS IN:		
		log. GNP	*BDEF*	*log. M*
		Interest Rate Differential (r^D)		
1	0.8497	0.0896	2.8712	8.3882
2	1.5984	4.1812	3.4602	9.6026
3	2.0025	10.4350	2.9521	9.9906
4	2.2698	12.5158	2.6492	9.0876
5	2.4427	13.8275	3.1736	8.5445
10	4.6855	17.1149	18.7632	15.3778
15	5.5776	19.4526	16.8002	17.00887
20	5.7855	19.0312	16.7319	17.1037
		Real Exchange Rate (e)		
1	21.8229	21.2781	0.0263	0.0057
2	69.3323	26.6466	1.2885	1.4202
3	146.4173	30.2225	1.8197	1.6750
4	261.2878	20.3387	4.6163	1.2617
5	375.5456	30.5192	8.4197	0.9930
10	594.9891	31.5615	16.1630	0.8942
15	724.8856	26.7346	26.6697	2.0221
20	979.2268	26.4127	31.2940	2.5778
		Current Account Balance (ca)		
1	2.6709	1.1434	0.9457	0.4083
2	3.5815	2.6096	1.4101	1.7047
3	4.6371	2.0156	2.9971	7.0486
4	5.3754	2.6100	5.1248	10.8263
5	7.1931	3.2635	6.5226	15.6079
10	14.0378	8.9799	8.9799	17.7991
15	26.1530	20.6530	20.6529	12.3564
20	32.5858	18.2228	18.2228	13.7549

TABLE 6.2(contd): **Percentage of *k*-step-ahead Forecast Error Variances of the Interest Rate Differential (r^D), the Real Exchange Rate (*e*) and the Current Account Balance (*ca*) accounted for by shocks to *log. GNP*, Budgetary Deficit (*BDEF*) and *log. M***

C) SWEDEN

k	VARIANCE OF k-STEP-AHEAD FORECAST ERROR	% VARIANCE IN *k*-STEP-AHEAD FORECAST ERROR EXPLAINED BY ORTHOGONALIZED INNOVATIONS IN:		
		log. GNP	*BDEF*	*log. M*
		Interest Rate Differential (r^D)		
1	1.3957	0.5168	0.8921	28.1140
2	1.9505	3.8756	0.9150	25.0288
3	2.6916	3.0716	1.4207	25.9616
4	3.7547	2.7117	1.1174	38.5465
5	4.5916	2.2456	4.1078	45.3188
10	6.5962	6.5211	7.2035	44.0285
15	7.1444	7.2566	8.1383	41.8916
20	8.0168	7.3622	8.0292	43.1938
		Real Exchange Rate (*e*)		
1	0.0369	1.0288	4.7243	0.4535
2	0.1521	1.0088	6.8020	0.1205
3	0.2868	0.6049	7.9914	1.0618
4	0.4299	0.4127	7.8957	4.5196
5	0.5854	0.3615	7.3511	12.1196
10	1.2076	0.4551	7.4029	43.2369
15	1.5324	0.9931	8.6436	45.2157
20	1.9077	7.2493	10.0680	39.3854
		Current Account Balance (*ca*)		
1	3.2328	0.4330	0.4798	8.3496
2	3.5276	0.6386	5.9162	8.0433
3	3.9888	0.6106	8.7970	9.0531
4	4.9559	11.6496	10.8244	8.6265
5	5.8235	13.9489	15.4181	7.3512
10	9.5883	10.8296	10.7913	22.8640
15	12.2479	17.1623	9.4088	26.9157
20	13.3758	18.0336	9.0048	27.7471

TABLE 6.2(contd): **Percentage of *k*-step-ahead Forecast Error Variances of the Interest Rate Differential (r^D), the Real Exchange Rate (*e*) and the Current Account Balance (*ca*) accounted for by shocks to *log. GNP*, Budgetary Deficit (*BDEF*) and *log. M***

D) UK

k	VARIANCE OF k-STEP-AHEAD FORECAST ERROR	% VARIANCE IN *k*-STEP-AHEAD FORECAST ERROR EXPLAINED BY ORTHOGONALIZED INNOVATIONS IN:		
		log. GNP	*BDEF*	*log. M*
		Interest Rate Differential (r^D)		
1	1.4320	0.8523	7.6855	5.3220
2	2.7632	0.8183	7.1132	5.6460
3	3.3595	2.2666	6.4549	4.5208
4	3.7380	4.0765	8.0626	5.3269
5	4.1964	4.0412	11.3441	8.0536
10	6.3333	5.3474	9.4538	13.4222
15	6.8079	6.0162	9.1990	13.8832
20	7.2339	9.7705	8.7194	13.9264
		Real Exchange Rate (*e*)		
1	0.0001	1.1393	0.0848	0.1376
2	0.0004	2.5612	0.0952	0.2114
3	0.0007	1.7201	0.1595	1.3771
4	0.0111	1.1649	0.1234	5.5533
5	0.0149	0.9076	0.1291	10.5961
10	0.0259	0.7325	0.5734	19.6780
15	0.0340	1.6111	2.1931	19.2513
20	0.0466	3.8707	1.9655	29.6442
		Current Account Balance (*ca*)		
1	1.5274	5.0786	2.4904	3.1474
2	2.0655	11.5006	2.0779	12.7281
3	2.5844	16.6653	4.8926	12.0904
4	3.3164	21.3114	4.2701	13.1426
5	3.8730	23.6938	3.9090	15.2026
10	5.4714	21.7853	3.1024	13.8901
15	5.8274	22.1674	2.9557	13.4020
20	6.0885	23.5856	2.8684	13.7704

Summary in Dutch

De ondergang van het Bretton Woods systeem van vaste wisselkoersen in de eerste helft van de jaren zeventig vormde het begin van talrijke theoretische en empirische onderzoeken ter verklaring van de waargenomen volatiliteit van wisselkoersen en veranderingen in de lopende rekening van de betalingsbalans. Ook in de eerste helft van de jaren tachtig werd de wereldeconomie geplaagd door omvangrijke en ongecoördineerde innovaties in monetaire en fiscale variabelen. Deze omvangrijke en ongecoördineerde innovaties leidde tot hoge en sterk fluctuerende reële rentes en wisselkoersonevenwichtigheden. Bij een onevenwichtige wisselkoers wijkt de reële wisselkoers sterk af van het lange termijn evenwichtige pad dat wordt bepaald door relatieve koopkrachtpariteit. Een groot aantal artikelen over wisselkoersen binnen de internationale financieringstheorie is een exponent van bovengenoemde ontwikkelingen. Prominente onderzoekers op dit gebied zijn bijvoorbeeld Paul Krugman, Michael Mussa, Rudiger Dornbusch, William Branson, Alan Stockman, John Huizinga, Paul de Grauwe, Ronald MacDonald en Mark Taylor. Ondanks het feit dat er veel onderzoek is gedaan naar wisselkoersen binnen de internationale financieringstheorie, bestaat er naast het fundamentele onderzoek naar nominale en reële determinanten van wisselkoersbewegingen, nog voldoende ruimte voor onderzoek naar oplossingen voor vraagstukken veroorzaakt door verschillen in de werking van de valutamarkt in de theorie en de praktijk.

Dit proefschrift vult de empirische internationale financieringstheorie over wisselkoersen op de volgende manier aan. Dit proefschrift introduceert en maakt gebruik van de nieuwste empirische methoden om schokken in produktiviteit en monetaire beleid te identificeren. De geïdentificeerde schokken worden vervolgens gebruikt ter verklaring van waargenomen fluctuaties in wisselkoersen en de lopende rekening. De studies in dit proefschrift zijn een mengeling van zowel de internationale monetaire theorie als ook de moderne tijdreeksanalyse teneinde de aspecten van het internationale monetaire systeem te doorgronden. Deze aspecten betreffen het effect van produktiviteits- en beleidsschokken op kapitaalstromen en bilaterale wisselkoersfluctutaties.

Hoofdstuk 1 - getiteld "Inleiding en samenvatting" - introduceert de verscheidene empirische onderwerpen die behandeld worden in de daaropvolgende hoofdstukken van dit proefschrift en geeft een korte samenvatting van de empirische bevindingen in elk hoofdstuk. Hoofdstuk 1 verklaart belangrijke empirische concepten zoals "common stochastic trends" en de identificatie van monetaire beleidsschokken.

Hoofdstuk 2 geeft een overzicht van de onderliggende fundamentele wisselkoerstheorieën. Tevens worden in dit hoofdstuk de begrippen koopkrachtpariteit en ongedekte rentepariteit toegelicht.

De hoofdstukken 3 en 4 in het eerste deel van dit proefschrift passen de "King-Plosser-Stock-Watson common stochastic trends approach" toe teneinde de door stochastische schokken veroorzaakte fluctuaties in investeringen, lopende rekening en bilaterale wisselkoersen te analyseren.

Hoofdstuk 3 onderkent het feit dat reële en nominale wisselkoersen worden gedreven door nominale als ook door reële factoren. De reële factoren waar men hierbij onmiddellijk aan denkt zijn produktiviteitsschokken. Dit hoofdstuk maakt gebruik van de "common trends approach" teneinde empirisch de uitwerking van schokken op die trends van wisselkoersbewegingen en veranderingen in de lopende rekening te analyseren. Uit de vastgelegde innovaties blijkt dat binnenlandse produktiviteitsschokken een meer significante uitwerking op wisselkoersfluctuaties en lopende rekening fluctuaties hebben dan nominale schokken. Deze bevindingen worden zowel voor de korte als de lange termijn gevonden.

Onze bevindingen, zoals die staan beschreven in hoofdstuk 4, bewerkstelligen niet alleen dezelfde resultaten met behulp van de common trends approach als Glick en Rogoff (1995) die de Solow parameter gebruiken ter benadering van produktiviteitsschokken, maar biedt de onderzoeker ook meer informatie over de dynamiek van de effecten van produktiviteitsschokken (die zijn geïdentificeerd) op veranderingen in de lopende rekening en daaraan gerelateerde kapitaalstromen. Bovenstaande benadering gebruik ik om de effecten van idiosyncratische (land-specifieke) en mondiale produktiviteitsschokken op de lopende rekening te analyseren. Mijn bevindingen stemmen overeen met wat theoretisch verwacht mag worden. Positieve idiosyncratische produktiviteitsschokken leiden op lange termijn tot een significante verslechtering van de lopende rekening van de betalingsbalans. Desbetreffende schokken vormen de belangrijkste verklaring voor de variantie van de voorspelfout inzake de lopende rekening en binnenlandse investeringen. Mondiale schokken daarentegen zijn niet geschikt om veranderingen in de lopende rekening te verklaren aangezien zij een gelijke uitwerking op alle landen hebben.

Hoofdstuk 5 van dit proefschrift presenteert een identificatie-schema voor monetaire politiek om twee vraagstukken in de literatuur te helpen oplossen. Deze vraagstukken zijn de *wisselkoers- puzzel* en de *termijndisagio-puzzel*. Ongedekte rentepariteit vereist dat een daling van de binnenlandse korte rente (ten opzichte van de buitenlandse korte rente) in eerste instantie gepaard gaat met een depreciatie van de binnenlandse valuta en in tweede instantie wordt gevolgd door een appreciatie van de binnenlandse valuta opdat marktpartijen bereid blijven binnenlandse activa in hun portefeuille te houden. Empirische bevindingen die niet overeenkomen met deze vereiste resulteren in de *termijndisagio-puzzel*. De *wisselkoers puzzel* is de tendens van de binnenlandse valuta (dat wil zeggen de valuta van G7 landen maar niet de Amerikaanse dollar) om in tegenspraak met de theorie te depreciëren ten opzichte van de Amerikaanse dollar, na een negatieve binnenlandse monetaire schok. De bevindingen in dit proefschrift bevestigen eerdere bevindingen uit de literatuur waarbij de *wisselkoers-puzzel* wordt geassocieerd met een volledig recursief identificatie schema terwijl de *termijndisagio-puzzel* het gevolg is van identificatie schema's voor monetaire politiek die of er niet in slagen de werkelijke operationele procedures van de monetaire autoriteiten van de landen in de steekproef weer te geven of die niet de "feedback" tussen wisselkoersbewegingen en de houding van de buitenlandse monetaire politiek incorporeren. Dit hoofdstuk incorporeert niet alleen de feedback van de wisselkoers op de monetaire politiek maar ook de interactie tussen binnen- en buitenlandse schokken in monetaire beleid. Met behulp van deze aanpak tracht ik empirisch het feitelijk proces van beleid te implementeren waarbij monetaire autoriteiten hun politiek gebruiken om wisselkoers en rente bewegingen glad te strijken. Het geïdentificeerde vector-autoregressieve (VAR) schema dat wordt gebruikt om dit soort monetair beleid te implementeren waarbij beleidsmakers hun beleid baseren op een monetaire condities index (MCI) blijkt erg plausibel te zijn in termen van theoretische voorspellingen van de korte termijn reactie van de rente, de wisselkoers en niet-monetaire variabelen op zowel binnen- en buitenlandse monetaire beleidsschokken. Een MCI is een lineaire combinatie van zulke monetaire variabelen zoals daar zijn korte termijn rentes, inflatie en bilaterale wisselkoersen die monetaire autoriteiten trachten te beïnvloeden. De geschatte dynamische reacties tonen verder aan dat inderdaad het geïdentificeerde VAR schema dat beleidsinterdependentie en wisselkoers feedback incorporeert, de puzzels kunnen helpen oplossen. In het bijzonder, in de meeste onderzochte landen leidt een monetaire inkrimping op korte termijn tot een appreciatie van de binnenlandse valuta ten opzichte van de Amerikaanse dollar gevolgd door een depreciatie van de binnenlandse valuta binnen één tot tien maanden na de monetaire schok.

Hoofdstuk 6 onderzoekt het effect van een expansieve fiscale schok op fluctuaties in de bilaterale wisselkoers en de lopende rekening. De empirische vernieuwing van dit hoofdstuk betreft het gebruik van een spaarzame combinatie van cointegratie en vector-autoregressie. Vele centrale bankiers - zowel Alan Greenspan (voorzitter van de Amerikaanse Federal Reserve) als Hans Tietmeyer (president van de Duitse Bundesbank) - geloven dat een verlaging van het begrotingstekort de externe waarde van de binnenlandse valuta vergroot ondanks het feit dat internationale macroeconomen het niet eens zijn over de richting van de wisselkoers na een verlaging van het begrotingstekort. Hoewel, binnen de context van algemene modellen van de wisselkoers zijn verlagingen van het begrotingstekort onontkoombaar geassocieerd met veranderingen in het netto buitenlands actief, veranderingen in waargenomen risico's van binnenlandse actieva (die de grootte van de gewenste risicopremie beïnvloedt die nodig is om beleggers te bewegen binnenlandse activa in hun portefeuille te blijven houden) en in verwachtingen omtrent de toekomstige waarde van de binnenlandse valuta. Veranderingen in het begrotingstekort hebben zowel korte als lange termijn gevolgen voor de externe waarde van de binnenlandse valuta. Op korte termijn zal een stijging van het begrotingstekort een plotselinge appreciatie van de binnenlandse valuta veroorzaken (de binnenlandse rente stijgt), dat vervolgens een handelstekort creëert en de lopende rekening verslechtert (kapitaalimport vanuit het buitenland). Deze kapitaalimport gaat door totdat de risicopremie (het toegenomen begrotingstekort verhoogt het waargenomen risico door beleggers van binnenlandse actieva) voldoende is toegenomen om verwachte binnenlandse en buitenlandse opbrengstvoeten te egaliseren. Dus op lange termijn geldt ongedekte rentepariteit weer en gaan positieve fiscale innovaties gepaard met lange termijn depreciatie van de binnenlandse valuta.

Dit proefschrift bespreekt kwesties die zowel voor de macroeconoom als de beleidsmaker van belang zijn. Prominent in dit proefschrift is de zoektocht naar een plausibel monetair identificatie schema dat in staat is de complexiteit van monetaire politiek in een open economie weer te geven. Hieraan verbonden is de zoektocht naar een oplossing voor de *termijndisagio-puzzel*. De zoektocht naar een empirisch plausibel monetair identificatie schema dat de *termijndisagio- puzzel* helpt op te lossen, blijft een zaak van belang die vele onderzoekers met mij nog de komende jaren zullen bezighouden.

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The demise of the Bretton Woods system of pegged exchange rates in the early 1970s gave birth to a number of theoretical and empirical research aimed at explaining the observed volatility of exchange rates and fluctuations in the current accounts of nations. Also, the early 1980s witnessed the world economy plagued by large and uncoordinated innovations in monetary and fiscal variables as well as by high and volatile real interest rates and exchange rate misalignments - i.e. the tendency for the real exchange rate movements to diverge from the long run equilibrium path as measured by relative purchasing power parity. The juxtaposition of these events led to a surge of scientific investigations which up to now, however, have not resolved the existing puzzles in a satisfactory way. This dissertation contributes to this literature on empirical international finance by introducing and using new empirical methods in identifying productivity shocks and monetary policy shocks. The shocks so identified are then used in explaining observed exchange rate and current account fluctuations. The rationale of the studies conducted is to blend international monetary theory with modern time series analysis in an attempt to contribute to our understanding of the current international monetary system. In particular, this dissertation presents empirical analysis on the effects of productivity and policy shocks on capital flows and bilateral exchange rate fluctuations using modern time series methods, and takes a step towards a deeper understanding of the obseerved phenomena.

ISBN 90-5668-024-2

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